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**Earnings management or forecast guidance to meet analyst expectations?**

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# Earnings management or forecast guidance to meet analyst expectations?

**Vasiliki E. Athanasakou, Norman C. Strong and Martin Walker***

**Abstract**—We examine whether UK firms engage in earnings management or forecast guidance to ensure that their reported earnings meet analyst earnings expectations. We explore two earnings management mechanisms: (a) positive abnormal working capital accruals; and (b) classification shifting of core expenses to non-recurring items. We find no evidence of a positive association between income-increasing, abnormal working capital accruals and the probability of meeting analyst forecasts. Instead we find evidence consistent with a subset of larger firms shifting small core expenses to other non-recurring items to just hit analyst expectations with core earnings. We also find that the probability of meeting analyst expectations increases with downward-guided forecasts. Overall our results suggest that UK firms are more likely to engage in earnings forecast guidance or, for a subset of larger firms, in classification shifting rather than in accruals management to avoid negative earnings surprises.

**Keywords:** meeting analyst expectations, abnormal accruals, earnings forecast guidance, classification shifting

## 1. Introduction

Survey-based evidence in the US (Graham et al., 2005) and the UK (Choi et al., 2006) shows that meeting analyst expectations is a fundamental earnings target. Severe stock market reactions to negative earnings surprises and a market reward to positive earnings surprises give managers strong incentives to walk down analyst earnings forecasts in order to increase the probability of hitting the final forecast (earnings forecast guidance) or to use their discretion over reported earnings to meet expectations (earnings management). In this paper, we examine whether UK firms use earnings management or forecast guidance to meet analyst expectations. In addition to the practice of accruals management, we examine a recently explored earnings management mechanism: inflating core earnings through classification shifting of core expenses to income-increasing (negative), non-recurring items. As both managers and analysts exclude non-recurring items from core earnings, firms may engage in classification shifting of recurring losses or expenses to inflate core earnings and meet analyst' expectations. Examining the practice of classification shifting by UK firms is of special interest, as over our study period FRS 3 *Reporting Financial Performance*, which was in force for UK firms from 1993 until the adoption of International Financial Reporting Standards in 2005, required firms to report net income per share, but allowed them to distinguish between core and transitory earnings by exercising discretion in classifying non-recurring items. In line with FRS 3's approach, IAS 1 *Presentation of Financial Statements* requires a clear distinction between core and exceptional income components, allowing firms to disclose material items of an exceptional nature separately in the income statement.

Prior evidence does not explore fully the link between the UK regulatory framework for reporting financial performance and the mechanisms UK firms use to meet expectations. In their survey of investment professionals and financial managers, Choi et al. (2006) report a consensus view that the general quality of earnings improved post-FRS 3 and that, while earnings forecast guidance might be a widespread phenomenon, firms are now less likely to use discretionary accounting choices to meet analyst expectations. At the same time, the survey reveals concern over the potential manipulation of non-recurring items despite the increased transparency requirements of FRS 3. The evidence of Choi et al. (2005) that a substantial proportion of the 500 largest UK listed non-financial firms (over 70%) exploited the option to disclose alternative earnings per share (EPS) on core earnings supports this concern. Athanasakou et al. (2007) lend further support to this concern by documenting an overall increase in the practice of income

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smoothing using classifications of non-recurring items post-FRS 3. Even though Peasnell et al. (2000a) and Gore et al. (2007) provide preliminary evidence consistent with UK firms using non-recurring items pre-FRS 3 to hit earnings benchmarks, they do not find similar evidence for the post-FRS 3 period. Therefore, to date there is no direct evidence on whether UK firms use classifications of non-recurring items post-FRS 3 to hit analyst forecasts.

In view of this discussion, we undertake an archive-based examination of the use of accruals management, classification shifting and earnings forecast guidance to meet analyst expectations in the post-FRS 3 period, 1994–2002. We constrain our sample period to 2002 due to data unavailability for key variables in our study post-2002. We use logistic analysis to examine the association of income-increasing, discretionary accruals and downward-guided forecasts with the probability of meeting analyst forecasts. As a proxy for discretionary accruals, we use abnormal working capital accruals (AWCAs) estimated using the modified Jones model including lagged return on assets (ROA) to control for operating performance (Kothari et al., 2005). Consistent with prior research (Burgstahler and Dichev, 1997; Dechow et al., 2000; Peasnell et al., 2000a; Das and Zhang, 2003), we focus on working capital accruals instead of total accruals, as changes in working capital offer a more flexible mechanism to meet earnings benchmarks than non-current accruals (e.g. depreciation, amortisation, impairments). We measure earnings forecast guidance by comparing an estimate of the expected analyst forecast (Matsumoto, 2002) with the last forecast for the year made before the release of the earnings announcement. Our findings show no evidence that firms use income-increasing AWCAs to meet analyst forecasts. Instead, we find a significant association between earnings forecast guidance and the probability of hitting the target. The relation persists after controlling for other incentives.

To examine whether UK managers shift core expenses to non-recurring items to meet analyst expectations we test the association between an estimate of unexpected core earnings (McVay, 2006) and non-recurring items, focusing on cases where classifications of non-recurring items allow managers to just hit analyst forecasts. We decompose total non-recurring items into non-operating exceptional items and other non-recurring items (i.e. operating exceptionals and other value-irrelevant items). We further distinguish between small and large non-recurring items. Other and small non-recurring items provide more latitude for classification shifting because they are less visible and less likely to depend on the occurrence of specific events (e.g. structural events such as mergers, acquisitions, restructurings and divestitures). Our results show a significant association between unexpected core profits and small income-increasing other non-recurring items for larger firms that would have just missed the analyst forecast without these classifications. For these firms, comprising around 10% of the sample,[1] the unexpected rise in core earnings appears to reverse in the subsequent accounting period when small core expenses recur, causing a decline in profitability. Additional analysis shows that while for the entire sample there is no association between non-recurring items and future operating performance, small other non-recurring items predict substantial operating cash outflows three years ahead for firms that just hit the analyst target with these classifications. Taken together, this evidence is consistent with a subset of larger UK firms engaging in classification shifting of small core expenses to other non-recurring items to meet analyst forecasts.

Our study provides information for accounting standard setters and contributes to the earnings management literature in several ways. First, we shed light on the earnings game in the UK by providing insights into the mechanisms UK firms use to meet analyst expectations and their incentives (e.g. relating to other earnings benchmarks, the value-relevance of earnings, litigation risk, growth and size). We find that post-FRS 3 UK firms did not use abnormal working capital accruals to meet analyst expectations. This corroborates survey evidence of Graham et al. (2005) and Choi et al. (2006) that managers are generally less likely to employ discretionary accounting adjustments to hit analyst forecasts. Also, using a proxy that captures both formal guidance (e.g. public disclosures) and informal guidance (e.g. through private conversations), we show that UK firms guide analyst forecasts down to meet analyst expectations. This evidence is consistent with the views of investment professionals and financial managers that earnings forecast guidance is a widespread practice in the UK (Choi et al., 2006).

Second, this is the first study to explicitly examine the possibility that UK firms shifted core expenses to non-recurring items post-FRS 3 to increase core earnings and hit the final forecast. McVay (2006) examines classification shifting by US firms using special items. Our evidence shows that even though classification shifting is not a common practice, managers of some larger firms may shift small core expenses to operating exceptional and other non-recurring items to inflate core earnings and hit the final forecast. This evidence justifies concerns over potential misclassifications

[1] These are 121 large firms with average total assets of £1,043m, average total sales of £1,426m and average market capitalisation of £1,576m.

of non-recurring items to meet earnings targets, despite the increased transparency requirements of FRS 3.

Third, our study adds to recent research examining adjusted earnings reporting in the UK (Walker and Louvari, 2003; Choi et al., 2005; Choi et al., 2007). Over the entire sample we find no association between non-recurring items and future operating performance. Consistent with prior evidence, these results suggest that, on average, UK firms classify transitory income components below core earnings. The only evidence consistent with misclassified non-recurring items relates to the subset of larger firms that would have just missed the final forecast without classifications of small other non-recurring items.

## 2. Prior research

### *2.1. US evidence*

In their survey of US chief financial officers, Graham et al. (2005) find that analyst expectations and prior earnings are managers' two most important earnings benchmarks. Strong stock market reactions to small negative earnings surprises suggest that investors view a missed target as signalling a substantial decline in underlying performance. Conversely, investors view a zero or positive earnings surprise as evidence of a well-managed firm, able to both predict and deliver future earnings. Archival research corroborates this evidence. Skinner and Sloan (2002) and Brown (2003) document adverse market reactions for firms that just fail to meet expectations. Bartov et al. (2002) and Kasznik and McNichols (2002) find that firms that meet analyst expectations earn a market reward that is positively associated with future performance. Doyle et al. (2006) show that this market reward is robust to other risk factors and previously documented market anomalies that affect variation in returns. Comparing the market penalty for missing analyst expectations to the premium for achieving analyst expectations, Lopez and Rees (2002) find that the former is significantly higher than the latter and that firms that consistently achieve analyst forecasts have a significantly higher earnings response coefficient. Measuring stock sensitivity by outstanding stock recommendations, Abarbanell and Lehavy (2003b) find that firms rated with a buy recommendation are more likely to inflate earnings to meet or slightly exceed analyst expectations. In addition to capital market consequences, evidence of Graham et al. (2005) suggests that managerial concerns over job security and reputation offer a strong incentive to achieve analyst expectations.

The importance of analyst forecasts as an earnings target appears to have risen over the last decade. Even though in the initial evaluation of the hierarchy of earnings targets by Degeorge et al. (1999) analyst expectations had the lowest rank, Brown and Caylor (2005) find that from the mid-1990s managers sought to avoid negative earnings surprises more than to avoid reporting losses or earnings decreases. The authors rationalise the switch in the target hierarchy by reporting evidence of a significantly higher reward (penalty) for achieving (missing) analyst expectations than for achieving (missing) the other two earnings targets. They attribute the higher premium to temporal increases in the accuracy and the precision of analyst forecasts, analyst following, media attention on meeting or beating analyst expectations and the number of firms followed by analysts.

Given the strong incentives to meet analyst expectations, substantial research has focused on the methods firms use to hit targets. The most popular techniques examined are positive abnormal accruals and earnings guidance to walk down analyst expectations to managers' desired figure. Burgstahler and Dichev (1997) provide evidence of firms using working capital accruals to avoid losses or earnings declines. They observe high changes in working capital for earnings that fall just above target, giving rise to a discontinuity in their distribution. Payne and Robb (2000) find that firms with pre-managed earnings below analyst forecasts have greater positive abnormal accruals. Dechow et al. (2000) find that firms that just meet analyst forecasts have higher abnormal working capital accruals than firms that just miss the target, while Das and Zhang (2003) document that managers use working capital accruals to round up reported EPS to meet analyst forecasts. Abarbanell and Lehavy (2003a) find that abnormal accruals are the main source of both the tail and the middle asymmetry in the distribution of forecast errors, indicating the use of accruals to meet analyst expectations in the current period or to increase the likelihood of hitting the target in the future. Matsumoto (2002) documents a positive association between income-increasing, abnormal accruals and the likelihood of avoiding negative earnings surprises. This evidence on accruals management comes from US data predating the accounting scandals of the early 2000s (e.g. Enron, Worldcom) and the Sarbanes Oxley Act of 2002. Based on a subsequent period, Graham et al. (2005) report that chief financial officers of US firms are generally reluctant to use discretionary accounting adjustments to hit earnings targets. In the post-scandals period, Koh et al. (2008) find that the propensity of US managers to rely on income-increasing, abnormal accruals to meet analyst expectations has decreased.

Bartov et al. (2002) provide preliminary evidence on earnings forecast guidance. They observe that even though analyst forecasts made at the beginning of a period overestimate earnings on aver-

age (giving a negative forecast error), as the end of the period approaches, analyst optimism turns to pessimism (giving a positive earnings surprise). They also find that the proportion of negative forecast errors ending with zero or positive earnings surprises is greater than the proportion of positive or zero forecast errors ending with negative earnings surprises. Matsumoto (2002) extends these findings by documenting a significant association between a proxy for earnings forecast guidance and the likelihood of beating expectations. Similarly, Burgstahler and Eames (2006) find that in addition to managing abnormal accruals, firms manage analyst earnings forecasts downward to just meet analyst expectations. In the post-scandals period, Koh et al. (2008) document an increase in the tendency of US firms to rely on earnings guidance to meet analyst expectations.

Recent US studies examine a potentially new device to meet analyst expectations: classifications of non-recurring items. While evidence suggests that customised earnings (e.g. pro forma earnings or Street earnings[2]) are more informative and more persistent than GAAP earnings (Bradshaw and Sloan, 2002; Bhattacharya et al., 2003), the financial press and accounting regulators in the US claim that managers may be behaving opportunistically by removing value-relevant items to hit earnings benchmarks. Doyle et al. (2003) justify regulatory scepticism with evidence that the items firms exclude from Street earnings are negatively associated with firms' future operating performance and that the market does not reflect this information. McVay (2006) adds to these findings with evidence that US firms re-classify core expenses as negative special items. She documents a positive association between a measure of unexpected core earnings and income-increasing, special items and finds that the association is stronger for firms that would have just missed the analyst forecast without the reclassification. Lin et al. (2006) provide consistent evidence when examining mechanisms that US firms use to meet or beat analyst expectations. Their evidence suggests that US firms use earnings forecast guidance and classification shifting and to a limited extent abnormal accruals to achieve analyst expectations. On average they find that the use of earnings forecast guidance, classification shifting and income increasing, abnormal accruals increases the probability of meeting or beating analyst expectations by 9%, 10% and 5% respectively.

### 2.2. *UK evidence and regulatory framework*

Evidence in the UK starts with the use of working capital accruals to achieve earnings targets. Peasnell et al. (2000a) document that UK firms with negative pre-managed earnings levels and changes have positive mean AWCAs. This holds for the period before the regulatory changes of the early 1990s. The enforcement of FRS 3, subsequent standards issued by the Accounting Standards Board (ASB) and the Cadbury Report reflected a general shift by UK regulatory bodies towards increased transparency and enhanced governance to restrain managerial attempts to manipulate earnings. Peasnell et al. (2000a) find that, post-Cadbury, the increased level of governance restrained the use of income-increasing AWCAs to avoid losses or earnings declines. This evidence is consistent with the regulatory shift and casts doubt on whether UK firms still use income-increasing AWCAs to achieve earnings targets in the post-FRS 3 period. Gore et al. (2007) examine the distributional properties of reported and pre-managed earnings and find that AWCAs are the main source of discontinuity of earnings levels throughout the period 1989–1998, but to a lesser extent post-FRS 3. However, examining the distributions of pre-managed and reported earnings to document the use of AWCAs to hit targets is sensitive to measurement error in AWCA estimates and likely to lead to erroneous inferences. Prior research has established that most abnormal accrual models are misspecified for firms that experience extreme operating performance (Dechow et al., 1995; Kothari et al., 2005). This results in misclassifications of normal accruals as abnormal in periods when firms are highly profitable or unprofitable. In this case, the result that pre-managed earnings do not display a discontinuity around zero cannot be clearly attributed to the use of abnormal accruals to exceed the benchmark. The misclassification argument becomes even more critical in view of the evidence of Brown (2001) that the discontinuity of earnings targets around zero is mainly a characteristic of profitable firms. To address this methodological limitation, we re-examine the use of AWCAs to achieve analyst expectations using a research design that is less subject to measurement error in AWCA estimates.

A further important element of the ASB's intervention in the financial reporting choices of UK firms was FRS 3's provisions for non-recurring items and the option to disclose customised earnings measures. FRS 3 required firms to distinguish between operating and non-operating exceptionals. Non-operating exceptionals included profits or losses on sales or termination of operations, fundamental reorganisation or restructuring costs and profits or losses on the disposal of fixed assets.

[2] Street earnings is the usual term for the adjusted earnings figure that analyst tracking services such as I/B/E/S, Zacks and First Call report as actual. Examples of charges that US firms exclude from these adjusted earnings are restructuring charges, write downs and impairments, R&D expenditures, merger and acquisition costs, mandatory stock compensation expense, goodwill amortisation, and certain results of subsidiaries (Bradshaw and Sloan, 2002).

Firms had to disclose these items under separate headings after operating profit, while they could disclose exceptional items relating to operations through a note or on the face of the income statement. Further to enhancing their disclosure, FRS 3 widened the definition of exceptionals to include any items of exceptional size or nature, thereby enabling firms to classify exceptional items according to the nature of the firm's operations. FRS 3 also redefined discontinued operations and required firms to disclose results from these operations separately in the income statement. Even more important, FRS 3 allowed firms to disclose alternative EPS on other profit levels, enabling them to remove non-recurring items and highlight a measure of core earnings. Pope and Walker (1999) report evidence of firms using exceptional items to classify bad news earnings components, mainly in the form of write-offs of large transitory losses. Consistent with this evidence, Walker and Louvari (2003) argue that disclosures of alternative EPS post-FRS 3 reflected managerial perceptions of persistent earnings. Choi et al. (2005) substantiate this argument with evidence that alternative EPS is more value-relevant than net income in terms of earnings predictability and price–earnings and return–earnings associations. Choi et al. (2007) find that the items (both gains and losses) managers exclude from alternative earnings definitions are value-irrelevant. Athanasakou et al. (2007) report an increase in persistence of pre-exceptional earnings post-FRS 3, suggesting that FRS 3 enhanced the role of classificatory choices over exceptional items in identifying sustainable profitability. While this evidence is inconsistent with misclassifications of non-recurring items, the broader scope for classificatory choices under FRS 3 may also have yielded a new mechanism for UK firms to meet analyst expectations. Gore et al. (2007) examine the distributions of positive and negative extraordinary items by earnings surprise portfolios pre-FRS 3 and find evidence consistent with firms classifying core expenses as extraordinary items to just meet analyst expectations. However, they do not find similar evidence for exceptional items post-FRS 3. Possible reasons are that the authors do not explore the association between exceptional items and pre-exceptional earnings in the absence of earnings management, or the properties of different types of exceptional items post-FRS 3.

Brown and Higgins (2005) provide initial evidence of earnings forecast guidance by UK firms. The authors argue that in strong investor protection environments, characterised by common law and market orientation, managers have greater incentives to avoid negative earnings surprises and they are more likely to use earnings forecast guidance than to manage reported earnings due to tighter financial reporting regulation and less rigorous regulation of forecast guidance. Consistent with their arguments, the authors find that the UK and other strong investor protection countries (e.g. the US, Sweden) have higher frequencies of firms engaging in earnings forecast guidance (approximately 37%) than countries with weak investor protection (e.g. Italy, Korea, Turkey).[3] In line with this initial evidence, the survey by Choi et al. (2006) shows that investment professionals and financial managers view earnings forecast guidance as a prevalent practice in the UK.

Our study contributes to this literature by examining whether UK firms are indeed more likely to use earnings forecasts guidance than earnings management to meet analyst expectations. In addition to abnormal accruals, we study classification shifting as a potential mechanism to hit the target in view of the broader scope for classificatory choices under FRS 3. Our study is similar to that of Lin et al. (2006) insofar as we examine mechanisms to meet analyst expectations. Nonetheless, the distinctive characteristics of the UK financial reporting framework allows us to conduct a more thorough examination by considering the relevant restraints (opportunities) that it imposes (creates) on the use of mechanisms to meet analyst expectations. To this extent our examination is more intuitive and offers direct insights to accounting standard setters, policy makers and market participants. These insights should be of special interest to international regulators, as they currently seek to optimise the framework for reporting financial performance (IASB 2008). In addition to its broader implications, our study uses a more sophisticated research design to test for accruals management and classification shifting, allowing for more valid inferences on the use of these practices to achieve analyst expectations.

## 3. Research design

Our objective is to examine whether UK firms engage in accruals management, classification shifting or earnings forecast guidance to meet analyst expectations. To examine earnings or expectations management to meet analyst expectations, we test the association between the probability of hitting the target and an earnings or expectations management proxy. To this end, we estimate abnormal accruals, misclassified non-recurring items and downward-guided analyst forecasts. For non-recurring items it is impossible to estimate the part

[3] La Porta et al. (1998) and Brown and Higgins (2005) show that the UK has the highest investor protection ranking among 21 countries (including the US, Australia, Canada, France Germany, Hong Kong, Greece, Spain Sweden and Switzerland). The general shift by UK regulatory bodies towards increased transparency and enhanced governance has contributed to the enforced investor protection in the UK.

resulting from intentional misclassifications because these items are unexpected and therefore unpredictable. Accordingly, for classification shifting we use a separate research design.

### *3.1. Abnormal accruals to meet analyst expectations*

As the end of an accounting period nears, managers can observe the firm's underlying earnings as well as the analyst forecast. During this time they can estimate any shortfall from the consensus forecast and use income-increasing AWCAs to eliminate it. However, analyst forecasts change up to the announcement of the results. From the financial year-end to the announcement of the results, managers cannot implement working capital accrual choices involving timing of transactions (e.g. accelerating sales). Other accrual choices (deferrals, accrued expenses, provisions, etc.) need estimating and booking within the accounting system. The process involves time constraints that reduce flexibility in using accruals to meet forecasts. Kasznik and McNichols (2002) show that the market penalises firms that previously met analyst expectations but subsequently fail to do so. This means there is a cost to meeting expectations through positive abnormal accruals, as their reversal may prevent managers from meeting expectations in future accounting periods. The aggressive use of positive abnormal accruals to meet analyst expectations is also more likely to raise the suspicions of auditors and the board of directors, who tend to scrutinise income-increasing, discretionary, accounting choices.

To capture income-increasing earnings management through working capital accruals, we use an indicator of positive abnormal working capital accruals (*POSAWCA*). Appendix A defines all the variables in the study. We focus on AWCAs instead of abnormal total accruals as in Matsumoto (2002) and Lin et al. (2006) for three reasons. First, research shows that working capital accruals account for most of the variation in total accruals (Sloan, 1996; Subramanyam, 1996; Thomas and Zhang, 2000; Dechow and Dichev, 2002). Second, working capital accruals are more flexible than non-current accruals (e.g. depreciation, amortisation, impairments) due to their frequent occurrence and the higher degree of judgment involved in their estimation. Third, non-current accruals usually represent large, visible, one-off costs or losses (e.g. write-downs, provisions for restructuring costs, impairments, losses on disposal of assets). Managers are unlikely to use these items to meet analyst expectations as analysts forecast core profitability, i.e. earnings before exceptional and other-non-recurring items.[4]

We estimate AWCAs using the cross-sectional modified Jones model with lagged ROA as a control for operating performance as follows:[5]

$$\frac{WCA_{it}}{A_{i,t-1}} = \lambda_0 \frac{1}{A_{i,t-1}} + \lambda_1 \frac{\Delta CR_{it}}{A_{i,t-1}} + \lambda_2 ROA_{i,t-1} + e_{i,t} \tag{1}$$

$$\frac{AWCA_{it}}{A_{i,t-1}} = \frac{WCA_{it}}{A_{i,t-1}} - \left[ \frac{\hat{\lambda}_0}{A_{i,t-1}} + \hat{\lambda}_1 \frac{\Delta CR_{it}}{A_{i,t-1}} + \hat{\lambda}_2 ROA_{i,t-1} \right]$$

where $\Delta CR_{it}$ and $AWCA_{i,t}$ are change in revenue net of change in accounts receivable and AWCAs for firm *i* in period *t*, $ROA_{it-1}$ and $A_{t-1}$ are return on assets and total assets for firm *i* in period $t$–1, $\lambda_0$ and $\lambda_1$ are regression parameters and $\hat{\lambda}_0$ and $\hat{\lambda}_1$ are OLS coefficients. We estimate equation (1) cross-sectionally within industry–years to correct for changing economic conditions that might affect accruals independently of earnings management. Peasnell et al. (2000b) show that the cross-sectional modified Jones model captures relatively subtle instances of accruals management in the UK. We estimate the model for each industry–year with at least six observations to ensure sufficient data for parameter estimation. We use Datastream Level 6 industry classifications.

### *3.2. Earnings forecast guidance to meet analyst expectations*

If earnings fall short of analyst expectations, firms can guide these expectations down to avoid excessively optimistic forecasts and increase the probability of meeting analyst expectations. UK investment professionals and financial managers view earnings forecast guidance as a common practice (Choi et al., 2006). Similar to positive abnormal accruals, earnings forecast guidance to meet analyst expectations entails costs. If initial forecasts are excessively high, guiding forecasts down requires analysts to revise their earnings expectations, which is likely to result in a negative market reaction at the forecast revision date. Continuous downward forecast guidance to keep forecasts at an achievable level can result in a period of falling prices. For earnings forecast guidance to be beneficial, the cost of a negative earnings surprise must exceed the cost of lower stock prices due to downward forecast revisions. This is likely to be the case

[4] Analysts forecast earnings on a continuing operations basis, before discontinued operations and exceptional and non-operating items, to retain persistent income components that are important for security valuation (I/B/E/S Glossary 2000: 8).

[5] As managerial discretion is not observable, measurement error in abnormal accruals is inherent. Prior research that evaluates accrual models (Dechow et al., 1995; Healy, 1996; Young, 1999; Thomas and Zhang, 2000) concludes that most accrual models do not control adequately for operating performance, exceptional and non-operating events, and growth. To mitigate concern over the measurement error in the accruals model (equation 1) we add lagged *ROA* as a control for operating performance and focus on working capital accruals which are less likely to be affected by exceptional and non-operating events, compared to non-current accruals. We also include a measure of growth in our multivariate specification (see equation 5).

as market reactions to earnings surprises tend to be stronger than market reactions to forecast revisions (Bartov et al., 2002: 189).

To obtain a proxy for earnings forecast guidance, we follow Matsumoto (2002). Her methodology allows us to disentangle innate analyst forecast revisions in response to bad earnings news from excessive downward revisions of the final analyst forecast and to focus on the latter to capture firms that guide analyst forecast down. We derive an indicator of downward-guided forecasts (*DOWN*) by comparing the last earnings forecast before the release of the earnings announcement (*AF0*) to an estimate, based on stock returns, of what the forecast would have been in the absence of guidance (*EF*). We define the unexpected forecast (*UEF*) as *AF0* minus *EF* and set *DOWN* to 1 (0) when *UEF* is negative (positive).

To estimate *EF* we use information that is available to analysts in forming their earnings expectations. We first model the change in I/B/E/S actual EPS ($\Delta EPS$) scaled by lagged share price ($P_{i,t-1}$) as a function of the prior year's change in earnings scaled by lagged share price and cumulative excess returns over the current year (*CRET*)

$$\Delta EPS_{i,t}/P_{i,t-1} = a_{1,t} + a_{2,t}\,(\Delta EPS_{i,t-1}/P_{i,t-2}) + a_{3,t}CRET_{i,t} + e_{i,t} \quad (2)$$

*CRET* is excess (market-adjusted) return cumulated from the month following the year *t*-1 earnings announcement to the month of the year *t* earnings announcement. This variable captures additional value-relevant information available to analysts in estimating their forecasts. We estimate equation (2) cross-sectionally within industry–years (using OLS), similar to the method of estimating AWCAs. Using cross sectional estimation with the time dimension fixed and scaling by lagged price means that stationarity issues arising from the inclusion of a lagged dependent variable in equation (2) is not a cause for concern. To mitigate the effect of outliers on parameter estimation, we winsorise the top and bottom 0.5% of the variables.[6] To calculate expected change in earnings, we use parameter estimates of the prior year, so as to use only data that analysts could have obtained when forecasting earnings

$$E(\Delta EPS_{i,t}) = [\hat{a}_{1,t-1} + \hat{a}_{2,t-1}\,(\Delta EPS_{i,t-1}/P_{i,t-2}) + \hat{a}_{3,t-1}CRET_{i,t}] \times P_{i,t-1} \quad (3)$$

*EF* is lagged I/B/E/S actual EPS plus the expected change in earnings $E(\Delta EPS)$

$$EF_{i,t} = EPS_{i,t-1} + E(\Delta EPS_{i,t}) \quad (4)$$

After obtaining *EF*, we derive *DOWN* as follows

$$UEF_{i,t} = AF0_{i,t} - EF_{i,t}$$

$$DOWN_{i,t} = \begin{cases} 1 & \text{if } UEF_{i,t} < 0 \\ 0 & \text{otherwise} \end{cases}$$

*3.3. Methodology and empirical predictions*

To examine whether the proxies for accruals management and earnings guidance capture the mechanisms UK firms use to meet analyst expectations we examine the relation between (a) the probability of meeting analyst forecasts; and (b) the indicators of positive AWCAs (*POSAWCA*) and of downward-guided analyst forecasts (*DOWN*), using contingency tables and multivariate models that control for other factors associated with meeting analyst expectations. We estimate the following logit regression of the probability that a firm meets or beats analyst forecasts at the earnings announcement date

$$\begin{aligned} Prob(MBE = 1|X) = F(&\beta_0 + \beta_1 POSAWCA_{i,t} \\ &+ \beta_2 DOWN_{i,t} + \beta_3 PROFIT_{i,t} + \beta_4 POS\Delta EARN_{i,t} \\ &+ \beta_5 VREARN_{i,t} + \beta_6 GROWTH_{i,t} + \beta_7 LIT_{i,t} \\ &+ \beta_8 INDPROD_{i,t} + \beta_9 SIZE_{i,t} + \beta_{10} MBE_{i,t-1} + u_{i,t}) \end{aligned} \quad (5)$$

where

$$F(\beta' X) = \frac{e^{\beta' X}}{1 + e^{\beta' X}}$$

*MBE* equals 1 if the earnings surprise is zero or positive, and 0 otherwise. The earnings surprise is actual EPS minus the latest analyst forecast made prior to the earnings announcement date, both from I/B/E/S. As prior UK evidence casts doubt on whether UK firms used AWCAs to meet analyst expectations post-FRS 3, the sign of the coefficient on *POSAWCA* is an empirical question. If UK firms guide analyst forecasts down to meet analyst expectations, consistent with the survey evidence of Choi et al. (2006), we expect the coefficient on *DOWN* to be positive.

In addition to meeting analyst expectations, the prior literature establishes two other earnings targets: profits and positive earnings changes. Degeorge et al. (1999) were the first to address the hierarchy of earnings targets. Based on conditional distributions of the three earnings benchmarks, they infer a 'pecking order', with profit first, prior year earnings second and analyst forecasts third. As they find no evidence of a discontinuity around zero in the distributions of forecast errors for firms reporting losses and earnings decreases, they conclude that meeting analyst expectations matters only if both other targets are met. Similarly, Brown (2001) finds that profitable firms make the most effort to meet analyst expectations and that the discontinuity around the earnings benchmark is not evident in loss-making firms. Even though Brown and Caylor (2005) suggest that since the mid-1990s managers

[6] We set the bottom 0.5% of the values of the variables equal to the value corresponding to the 0.5th percentile, and the top 0.5% of the values to the value corresponding to the 99.5th percentile.

have sought to avoid negative earnings surprises more than losses and earnings declines, Dopuch et al. (2003) find that there is an incremental market credibility premium for firms that meet analyst expectations in addition to meeting a time series benchmark. Their evidence suggests that firms have stronger incentives to meet analyst forecasts once they meet prior year earnings. In their survey, Graham et al. (2005) provide evidence consistent with the priority of meeting prior-year earnings over analyst expectations. As evidence on the priority of earnings targets is controversial and focuses on US firms, the association between the three earnings targets in the UK remains an empirical question. If there is a pecking order in meeting earnings targets and analyst forecasts is last in the order, we need to control for the targets of profits and earnings increases. Accordingly, we add a profit indicator, *PROFIT*, and an indicator of positive change in earnings, *POSΔEARN*. We expect the coefficients on *PROFIT* and *POSΔEARN* to be positive.

Matsumoto (2002) argues that the underlying reason for Brown's (2001) findings is the low value-relevance of the earnings of loss-making firms and not merely the sign of contemporaneous earnings. She argues that managers of firms with low value-relevance of earnings are less likely to be concerned about hitting analyst forecasts since investors put less emphasis on a poor indicator of performance. Accordingly, we add a proxy for the value-relevance of earnings (*VREARN*) and predict that firms with low value-relevance of earnings are less likely to meet analyst expectations. Therefore, we expect the coefficient on *VREARN* to be positive.

Skinner and Sloan (2002) find that the discontinuity in the distribution of earnings surprises around zero is stronger for higher-growth firms. Managers of such firms have stronger incentives to avoid negative earnings surprises because the market reaction to earnings announcements is more severe for these firms. Moreover, Dechow et al. (2000), show that firms with zero earnings surprises have, along with high market-to-book ratios, high levels of accruals. Therefore we control for *GROWTH* and expect firms with higher growth opportunities to have greater incentives to meet analyst expectations. Similar to Skinner and Sloan (2002), we measure growth using the market-to-book ratio. We expect the coefficient on *GROWTH* to be positive.

Firms with a higher risk of shareholder litigation have greater incentives to meet analyst expectations and reduce the possibility of shareholder lawsuits. To control for this effect Matsumoto (2002) includes an industry dummy (*LIT*) identifying firms in the high risk industries of biotechnology, computers, electronics and retailing. She argues that the industry indicator offers a better measure of ex-ante litigation risk, as it is unaffected by earnings management. Although we do not expect shareholder litigation risk to be as important for UK firms, we add *LIT* to control for any possible effect. To the extent firms in the high risk industries of biotechnology, computers, electronics and retailing have greater incentives to meet analyst expectations, we expect the coefficient on *LIT* to be positive.

We control for contemporaneous shocks to earnings since they are strongly related to analyst forecast errors. A firm's performance for the period, as well as general macroeconomic shocks, affect the probability of meeting analyst expectations. Positive earnings shocks are more likely to result in positive forecast errors than are negative earnings shocks. We control for this effect through *POSΔEARN*. To control for the impact of macroeconomic conditions on forecast errors, we include average annual growth in industrial production (*INDPROD*). Similar to Matsumoto (2002), we expect the frequency of meeting analyst expectations to increase with growth in industrial production, thus we expect the coefficient on *INDPROD* to be positive. Furthermore, as larger firms are more subject to public scrutiny their managers have stronger incentives to meet analyst expectations. Therefore we control for size and expect the coefficient on *SIZE* to be positive. Finally, Kasznik and McNichols (2002) find that firms that meet analyst expectations in consecutive years receive a market premium that is incremental to the higher future earnings that investors can rationally expect from these firms. This signifies the existence of a 'credibility' reward for consistently meeting expectations, giving greater incentives to meet analyst expectations in the current year if the firm met expectations in the previous year. We add $MBE_{t-1}$ to control for this effect and expect its coefficient to be positive.[7]

Doyle et al. (2004) argue that removing observations with large forecast errors and focusing around the zero earnings surprise benchmark gives a more powerful test of earnings or expectations management to hit the forecast. Therefore, we repeat equation (5) using JMBE instead of MBE, which captures cases where firms just meet analyst forecasts over cases where firms just miss the target, as follows:

$$\begin{aligned} Prob(JMBE=1|X) = F(&\beta_0 + \beta_1 POSAWCA_{i,t} \\ &+ \beta_2 DOWN_{i,t} + \beta_3 PROFIT_{i,t} + \beta_4 POS\Delta EARN_{i,t} \\ &+ \beta_5 VREARN_{i,t} + \beta_6 GROWTH_{i,t} + \beta_7 LIT_{i,t} \\ &+ \beta_8 INDPROD_{i,t} + \beta_9 SIZE_{i,t} + \beta_{10} MBE_{i,t-1} + u_{i,t}) \end{aligned} \tag{6}$$

[7] We also test a specification adding $MBE_{t-2}$ and $MBE_{t-3}$ (see footnote 15).

*JMBE* equals 1 if the earnings surprise (*SURP*) is in the interval £0.00 ≤ *SURP* < £0.02 and 0 if the earnings surprise is in the interval –£0.02 ≤ *SURP* < 0.[8] Using this specification we delete observations outside the interval –£0.02 ≤ *SURP* < £0.02.[9]

### 3.4. Classification shifting to meet analyst expectations

The provisions of FRS 3 broadened the scope for classificatory choices over non-recurring items. As both managers and analysts tend to exclude non-recurring items from core earnings, classification shifting of core expenses or losses to non-recurring items after the end of the accounting period and before the announcement of the results could increase the probability of hitting the final forecast. Unlike discretionary accounting choices, classification shifting does not flow through the accounting system, giving managers a more valuable timing option to meet the final forecast. In addition, classification shifting does not affect net income, thus limiting the scrutiny of auditors, outside monitors and regulators (Nelson et al., 2003; McVay, 2006).[10] Finally, classification shifting does not affect future period income, reducing substantially its cost as an earnings management device.

FRS 3's transparency requirements imposed an important cost of classification shifting on UK firms. FRS 3 required firms that reported alternative EPS to reconcile this to the basic figure, to disclose it consistently over time, to give it no greater prominence in the annual report than basic EPS and to explain the reasons for any alternative measure and its significance, adjacent to alternative EPS or through a note. Furthermore, for each exceptional item in the income statement, firms had to provide an adequate description to allow users to understand its nature. Transparency requirements add a constraint to attempts to shift core expenses below core earnings, as they increase the likelihood that investors detect misclassified non-recurring items. Although McVay (2006) does not explore potential changes in the practice of classification shifting by US firms after the Sarbanes Oxley Act of 2002, recent evidence suggests that the disclosure requirements of the Act had a significant impact on US managers' attempts to misclassify income components (Heflin and Hsu, 2008).

Even though transparency requirements restrain classification shifting, investors may be unable to distinguish cases of misclassified non-recurring items. This is especially true of operating exceptional items, as it is more difficult to determine their degree of persistence. In addition, while auditors may question classification manipulations involving more visible and specific non-operating exceptionals, they are less likely to challenge classifications of operating exceptionals. This is because the latter classifications rely more on managerial judgment and less on explicit definitions and disclosure requirements.

To examine whether UK firms reclassified core expenses as non-recurring items to meet analyst expectations in the post-FRS 3 period, we test the association between unexpected core earnings (levels and changes) and income-increasing, non-recurring items. Following McVay (2006), we derive measures of expected core earnings (*CE*) levels and changes ($\Delta CE$) as follows:

$$CE_{i,t} = \gamma_0 + \gamma_1 CE_{i,t-1} + \gamma_2 ATO_{i,t} + \gamma_3 WCA_{i,t-1} + \gamma_4 WCA_{i,t} + \gamma_5 \Delta SALES_{i,t} + \gamma_6 NEG\_\Delta SALES_{i,t} + u_{i,t} \tag{7}$$

$$\Delta CE_{i,t} = \gamma_0 + \gamma_1 CE_{i,t-1} + \gamma_2 \Delta CE_{i,t-1} + \gamma_3 \Delta ATO_{i,t} + \gamma_3 WCA_{i,t-1} + \gamma_4 WCA_{i,t} + \gamma_5 \Delta SALES_{i,t} + \gamma_6 NEG\_\Delta SALES_{i,t} + u_{i,t} \tag{8}$$

As a proxy for core earnings (CE) we use I/B/E/S actual EPS[11] multiplied by the weighted average

[8] To ensure equally sized intervals, we include –£0.02 in the interval of firms that just miss analyst expectations, but exclude £0.02 from the interval of firms that just meet or beat analyst expectations.

[9] Our core results remain when using a narrower *JMBE* interval –£0.01 ≤ *SURP* < £0.01.

[10] Auditors and outside monitors are more concerned with issues of unrecognised expenses, or abnormal variations in margins and other ratios, and are less likely to question the appropriate classification of expenses. Also firms reporting extraordinary and exceptional items tend to be declining in performance, so even though their core earnings might be higher than actual core earnings, their reported earnings may still be below prior-period and industry benchmarks.

[11] We choose I/B/E/S actual EPS as a proxy for core earnings for three reasons. First, compared to alternative EPS disclosed in the income statement, using I/B/E/S actual EPS allows us to expand the sample and run more powerful tests using a large panel of data instead of a small hand-collected sample. Second, I/B/E/S actual EPS is closer to analysts' definitions of earnings, which is important when calculating earnings surprises and our measure of *(J)MBE*. Bhattacharya et al. (2003) argue that to the extent there is a mismatch between the actual earnings and the forecast figure, there can be a severe error-in-variables problem in the earnings surprise measure. Third, compared with alternative adjusted earnings metrics available in Datastream, we expect I/B/E/S actual EPS to be a more accurate proxy of the firm's core earnings. This is because analyst tracking services adjust realised earnings by making exclusions for non-recurring items on what Gu and Chen (2004) refer to as a 'case by case' basis instead of the category by category basis that Datastream follows. This means that the tracking services treat transitory items selectively according to firm-specific characteristics, with the result that their exclusions are closer to those of managers. To assess the relation between alternative EPS, I/B/E/S actual EPS, and Datastream adjusted EPS we use data on alternative EPS disclosures for 1996 and 2001 for the 500 largest UK listed non-financial firms. We are grateful to Dr. Young-Soo Choi, Lancaster University, for providing this data. Descriptive analysis of this sample indicates that alternative EPS is most highly correlated with I/B/E/S actual EPS and that the mean and median differences between the two figures are not statistically significant.

number of shares (both unadjusted for splits)[12] and scaled by total sales. Lagged core earnings controls for earnings persistence over time. Given the close association between core earnings and profit margin, the asset turnover ratio (*ATO*) controls for the inverse relation with the profit margin, especially for firms with large income-increasing, non-recurring items (e.g. restructuring or reorganisation costs). Lagged working capital accruals ($WCA_{t-1}$) capture the information content of prior-year accruals for current-period income. Current working capital accruals ($WCA_t$) controls for extreme operating performance, as it is highly correlated with accrual levels. As excessive accruals could also reflect accruals management, controlling for $WCA_t$ is necessary to capture any excess profits associated solely with classification shifting. Finally, change in sales ($\Delta SALES$) controls for the effect of sales growth on fixed costs. As this effect differs between sales increases and decreases, we allow a separate coefficient for sales declines ($NEG\_\Delta SALES$). The model of change in *CE* (equation 8) is not obtained by merely differencing equation (7). Including both lagged *CE* and lagged $\Delta CE$ allows the degree of mean reversion to vary with the level of prior-year earnings.

We estimate equations (7) and (8) cross-sectionally within industry–years (using OLS). Unexpected core earnings (*UCE*) and unexpected change in core earnings ($U\Delta CE$) are the differences between reported core earnings and change in core earnings and their predicted values, which we derive based on the coefficients of equations (7) and (8). To examine whether classification shifting represents an earnings management tool to meet analyst expectations, we test the association between *UCE* and $U\Delta CE$ and income-increasing, non-recurring items for the entire sample and subsets of firms that are increasingly likely to use classification shifting to hit the target. In particular, similar to McVay (2006) we estimate the following two OLS regressions:

$$UCE_{i,t} = \delta_0 + \delta_1 TNRI_{i,t} + u_{i,t} \tag{9}$$

$$U\Delta CE_{i,t+1} = \delta_0 + \delta_1 TNRI_{i,t} + u_{i,t} \tag{10}$$

where TNRI is income-increasing, total non-recurring items defined as:

TNRI = I/B/E/S actual earnings – Net Income/Sales

and I/B/E/S actual earnings is the I/B/E/S reported actual EPS multiplied by the weighted average number of shares (both unadjusted for splits). Net income is earnings after extraordinary items. If I/B/E/S actual earnings is greater than net income, then total non-recurring items are income-increasing (through the exclusion of negative items). We estimate equations (9) and (10) for the entire sample and two subsets of firms: (a) firms with income increasing total non-recurring items ($TNRI > 0$); and (b) firms that would have just missed the analyst forecast had they not re-classified non-recurring items ($JUSTMET = 1$). JUSTMET equals 1 if the earnings surprise is from £0.00 to £0.02 per share, total non-recurring items are income-increasing and the earnings surprise minus total non-recurring items per share is negative, 0 otherwise. To the extent UK firms classification shift, unexpected core earnings increase with non-recurring items in year *t*, giving a positive $\delta_1$. If classification shifting serves as a mechanism to meet analyst expectations, the positive association between *TNRI* and *UCE* will be more profound for the two subsets of firms. While a positive association between *TNRI* and *UCE* is consistent with classification shifting, it is also consistent with efficiency gains resulting from disposals of unprofitable subsidiaries or from rationalising operations. To distinguish between these two competing hypotheses, we test whether the increase in *UCE* reverses in year *t*+1, as the core expenses firms misclassify in year *t* recur in year *t*+1, similar to McVay (2006). A positive $\delta_1$ in equation (9) and a negative $\delta_1$ in equation (10) are more consistent with classification shifting.

Choi et al. (2005) classify total non-recurring items that UK firms tend to exclude from core earnings into five groups: (a) non-operating exceptional items; (b) operating exceptional items; (c) other non-operating exceptional items; (d) charges relating to asset values; and (e) other non-recurring items. Non-operating exceptionals are usually large transitory items resulting from structural events. Their specific nature allows investors to assess their persistence with greater confidence. Moreover, the incidence and value of the finite number of events giving rise to such exceptionals constrain the extent to which managers can use them for classification shifting (Godfrey and Jones, 1999). The remaining items include: specified or unspecified exceptional operating costs; exceptional dividends, interest or taxation; other provisions; goodwill amortisation; amortisation of unspecified intangible assets; impairment diminution or write-off of goodwill; revaluation or impairment of fixed assets and discontinued operations; and acquisition, merger and demerger costs. These items offer greater latitude for classification shifting as their classification relies heavily on managerial judgment. To accommodate the different degree of clas-

[12] All data in the I/B/E/S detail file that are most commonly used in prior research appear on a split-adjusted basis to ensure that per share amounts are comparable over time. To calculate non-recurring items we need the historical figures for I/B/E/S actual EPS. Accordingly we 'unsplit' I/B/E/S actual EPS using the proper adjustment factors to derive the originally reported amounts.

sification discretion involved in the groups of non-recurring items, we extend the two equations of McVay (2006) by decomposing total non-recurring items (*TNRI*) into non-operating exceptional items (*NOEI*) (group (a)) and other non-recurring items (*ONRI*) (groups (b)–(e)) as follows

$$UCE_{i,t} = \delta_0 + \delta_{1a}NOEI_{i,t} + \delta_{1b}ONRI_{i,t} + u_{i,t} \qquad (11)$$

$$U\Delta CE_{i,t+1} = \delta_0 + \delta_{1a}NOEI_{i,t} + \delta_{1b}ONRI_{i,t} + u_{i,t} \qquad (12)$$

*NOEI* is non-operating exceptional items, i.e. profits or losses on the sale or termination of operations, costs of fundamental reorganisations or restructuring and profits or losses on the sale of fixed assets, adjusted for tax and minority interest and scaled by sales. We derive *ONRI* as the residual non-recurring items after deducting *NOEI* from *TNRI*. As other non-recurring items offer a more flexible device for classification shifting to meet analyst expectations we expect $\delta_{1b}$ to be positive.

Abarbanell and Lehavy (2002) find that when core earnings slightly exceed analyst forecasts while GAAP earnings fall slightly below, core earnings are more value-relevant than GAAP earnings. Therefore, they suggest that future research should isolate cases where classifications of small income-increasing non-recurring items lead to small positive earnings surprises. The magnitude of non-recurring items is an interesting property when considering classification shifting to meet analyst expectations. Large non-recurring items are highly visible and depend on the occurrence of specific (e.g. structural) events. Conversely, small non-recurring items are less debatable and offer a more flexible means of classification shifting, especially in cases where pre-managed earnings fall slightly short of the final analyst forecast. To incorporate this feature into our analysis, we repeat equations (9)–(12) on subsets of firms that just meet analyst expectations using small income-increasing, non-recurring items. We distinguish small non-recurring items similar to Doyle et al. (2004). *SmallTNRI* equals 1 if *TNRI* is positive and less than median *TNRI* (scaled by sales) for firms with positive *TNRI* in the same year.

As we use panel data, we need to control for cross-sectional dependence. To alleviate bias in the standard errors of equations (9)–(12) due to heteroskedasticity and cross-sectional dependence, we estimate standard errors clustered by year.

## 4. Sample selection

We collect data for all UK (dead and live) listed firms from Datastream for the period 1994 to 2002. We begin with 1994 because we focus on the period following the enforcement of FRS 3 (23 June 1993) and coverage of UK firms on I/B/E/S is limited for earlier years. Our sample period cut-off is 2002 due to data unavailability for non-operating exceptional items and other non-recurring items after the merger of Datastream with Worldscope in 2003.[13] We exclude financial firms because of their substantially different financial reporting environment and utilities because they are regulated and their earnings growth is typically more predictable. We also eliminate all accounting periods where there is a change in year-end, to ensure forecasts throughout the year refer to a 12-month accounting period. In line with Bartov et al. (2002), we keep observations with at least three individual earnings forecasts for the year to ensure there is an initial forecast, a reliable revision and a final forecast. The resulting sample with full data coverage in I/B/E/S is 1,397 firms and 6,199 observations. From this sample, consistent with McVay (2006), we delete observations with sales of less than £0.6m to avoid extreme outlying values of levels and changes in core earnings as sales are used as their deflator. Of the remaining observations (1,388 firms and 6,158 observations), we keep those where we can estimate unexpected core earnings. The final sample comprises 1,154 firms and 5,117 observations for tests of *MBE* and 1,033 firms and 3,609 observations for tests of *JMBE*. Tests of future unexpected changes in core earnings require one-year-ahead earnings, reducing the sample to 941 firms and 3,968 observations. To mitigate the effect of outliers in the data, we winsorise all variables at the 0.5% and 99.5% percentiles.

The first two columns of Table 1, Panel A report the annual distribution of observations for tests of *MBE*. The number of firms in each period represent on average 77% of all UK non-financial firms covered by I/B/E/S with data available for analyst forecasts and actual earnings (not tabulated). The third column of Table 1, Panel A reports the annual distribution of observations for tests of *JMBE*. The fraction of earnings surprises within the interval for measuring *JMBE* ($-£0.02 \leq SURP < £0.02$), is approximately 71% of the sample (3,609 out of 5,117) indicating a high concentration of UK firm–year observations around the earnings surprise benchmark.

[13] As Datastream recently retrieved data for accounting items up to 2005 (Datastream archive) we assessed the scope for extending our sample period post-2002. After inspecting the archive we concluded that the data was not reliable due to limited firm coverage and minor discrepancies with the old Datastream data for key variables in our study. Since these data limitations give rise to survivorship bias and errors in variables issues in the extended sample, we maintain the reported results on the original sample (1994–2002). As a robustness check we repeated our core tests on a sample extended to 2005 (pre-IFRS) using data from the archive. We obtain qualitatively similar results. The only exception pertains to tests of classification shifting for subsets of firms that are more likely to have engaged in this practice. As tests of classification shifting on a narrow subset of firms are especially prone to survivorship bias this result is not appropriate for valid inferences.

**Table 1**

**Panel A: Frequency of meeting analyst forecasts (*MBE* = 1), frequency of just meeting analyst forecasts (*JMBE* = 1), and negative forecast errors**

| | | | *MBE=1* | | *JMBE=1* | | | |
|---|---|---|---|---|---|---|---|---|
| | *Entire sample N* | *–£0.02≤ SURP<£0.02* | *N* | *Freq.(%)* | *N* | *Freq.(%)* | *Negative forecast errors (%)* | *Mean (median) forecast error* |
| All years | 5,117 | 3,609 | 3,086 | 60.31 | 2,346 | 65.00 | 60.80 | –3.44(–0.74) |
| 1994 | 503 | 388 | 350 | 69.58 | 272 | 70.10 | 51.89 | –2.26(–0.20) |
| 1995 | 562 | 461 | 363 | 64.59 | 308 | 66.81 | 56.05 | –2.42(–0.40) |
| 1996 | 585 | 461 | 378 | 64.62 | 305 | 66.16 | 60.34 | –2.80(–0.50) |
| 1997 | 638 | 504 | 437 | 68.50 | 354 | 70.24 | 53.45 | –2.03(–0.20) |
| 1998 | 649 | 489 | 431 | 66.41 | 338 | 69.12 | 56.70 | –2.88(–0.50) |
| 1999 | 595 | 407 | 382 | 64.20 | 266 | 65.36 | 57.31 | –2.62(–0.60) |
| 2000 | 525 | 318 | 275 | 52.38 | 192 | 60.38 | 61.14 | –4.78(–1.10) |
| 2001 | 548 | 315 | 236 | 43.07 | 169 | 53.65 | 75.55 | –6.08(–2.99) |
| 2002 | 512 | 266 | 234 | 45.70 | 142 | 70.10 | 74.80 | –5.65(–3.25) |

**Panel B: Real growth in UK GDP, the industrial production index and the index of total services**

| | *Real growth in GDP* | *Real growth in industrial production* | *Real growth in services* |
|---|---|---|---|
| All years | 0.031 | 0.009 | 0.019 |
| 1994 | 0.043 | 0.057 | – |
| 1995 | 0.030 | 0.007 | –0.011 |
| 1996 | 0.028 | 0.014 | 0.013 |
| 1997 | 0.031 | 0.000 | 0.018 |
| 1998 | 0.034 | 0.006 | 0.032 |
| 1999 | 0.030 | 0.025 | 0.027 |
| 2000 | 0.038 | 0.019 | 0.039 |
| 2001 | 0.024 | –0.047 | 0.021 |
| 2002 | 0.021 | –0.002 | 0.014 |

The sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms meeting the sample selection criteria. All observations outside the interval –£0.02 ≤ *SURP* < £0.02 are deleted when *JMBE* is used, resulting in 1,033 firms and 3,609 observations. *SURP* is the earnings surprise measured as the difference between I/B/E/S actual EPS and the latest analyst forecast made prior to the earnings announcement date. The number of observations with zero or positive earnings surprises (*MBE*=1) is 3,086. The number of observations with zero or small positive earnings surprises (*JMBE*=1) is 2,346. Forecast error is the difference between I/B/E/S actual EPS and the earliest forecast for the year (following the announcement of the previous year's earnings). We obtain the real growth in GDP and in industrial production from Datastream using the following codes: UKGDP%D (annual real change in UK GDP), UKCKYWA (UK industrial production: all production industries) and UKFVQQ (index of total services) and UKCONPRCF (UK RPI). Appendix A defines the remaining variables.

## 5. Results

### 5.1. Descriptive analysis

Columns 4–7 of Table 1, Panel A report the frequency of meeting analyst forecasts (*MBE*) and just meeting analyst forecasts (*JMBE*) across years. While Matsumoto (2002) documents a monotonic increase in *MBE* from 1990 to 1997, in the UK *MBE* does not change systematically from 1994 to 1997, but there is a steady decrease starting in 1998, becoming sharper in 2000 and 2001. Over the same period *JMBE* also falls. To explore the fall in *MBE*, the next two columns of Table 1, Panel A report the frequency of negative forecast errors and the mean and median forecast error, while Panel B reports the annual real growth in the UK GDP, industrial production and the index of total services. The frequency of negative forecast errors rises sharply in 2001, reaching 75.55%. Mean and median forecast errors are increasingly negative in 2000 and 2001. The UK economy appears to have experienced a mild recession during 2001–2002. Weak macroeconomic conditions are reflected in the falling rate of real growth in the UK GDP (from 0.038 in 2000 to 0.024 in 2001 and 0.021 in 2002) and the index of total services (from in 0.039 in 2000 to 0.021 in 2001 and 0.014 in 2002) and the fall in industrial production (–0.047 in 2001 and –0.002 in 2002). Overall, the fall in *MBE* overlaps a period of macroeconomic decline and a rise in the magnitude and frequency of negative forecast errors.

Panel A of Table 2 reports descriptive statistics for key variables. Mean *MBE* and *JMBE* for the entire panel are 60% and 65%. While the average use of positive AWCAs is 49%, 53% of forecasts are guided down.[14] Mean (median) core earnings as a percentage of sales are 1.9% (5%). Mean non-recurring items are 2.1% of total sales: 4.6% when they are income-increasing and –0.9% when they are income-decreasing (from untabulated analysis). A small part of non-recurring items, 0.3% of total sales on average, comprises non-operating exceptional items (*NOEI*), while the major part, 1.8% of total sales, is other non-recurring items (*ONRI*) (e.g. operating exceptionals and other non-operating and value-irrelevant items). Annual analysis (untabulated) shows a gradual rise in *TNRI* from 0.9% of sales in 1994 to 5.6% in 2002, due mainly to the growing magnitude of *ONRI*.

Panel B of Table 2 reports Pearson (Spearman) correlations between the key variables. While the frequency of achieving analyst expectations is positively correlated with the frequency of forecasts guided down (0.080, 0.080) and unexpected core earnings (0.088, 0.144), it is not significantly associated with reporting of positive AWCAs. Reporting of positive AWCAs is negatively associated with unexpected core earnings (–0.034) and income-increasing other non-recurring items (–0.032, –0.026). The frequency of forecasts guided down is also negatively correlated with unexpected core profits (–0.048, –0.105). This evidence is consistent with managers using earnings and expectations management mechanisms as substitutes. Finally, income-increasing other non-recurring items are positively correlated with unexpected core profits (0.033, 0.029) and negatively correlated with future changes in unexpected core earnings (–0.042). These results suggest that firms with guided forecasts and unexpected core profits arising from classifications of other non-recurring items are more likely to achieve analyst expectations.

### 5.2. Results on accruals management vs. forecast guidance to meet analyst expectations

#### 5.2.1. Contingency tables

Figure 1 shows the distribution of earnings surprises. We aggregate earnings surprises observations into equally sized intervals (bins). We set the size of each bin to 0.5p. As we concentrate on earnings management attempts to hit the zero earnings surprise benchmark, we aggregate all earnings surprises below –10p to the 20th bin below zero and all earnings surprises above 10p to the 20th bin above zero. Consistent with Gore et al. (2007: 132) we document a discontinuity at zero caused by the higher frequency of small positive compared with small negative surprises. This is consistent with either earnings management or forecast guidance to meet analyst expectations. Table 3, Panel A presents contingency tables examining the relation between *MBE* and indicators of positive AWCAs (*POSAWCA*) and of downward-guided analyst forecasts (*DOWN*). The results show no significant difference in *MBE* between firms with income-increasing (59.27%) versus income-decreasing or zero AWCAs (61.31%). In contrast, of firm–years where analyst forecasts are guided down (*DOWN* = 1) 64% meet analyst expectations, compared with firm–years with no downward guidance (*DOWN* = 0) when 56% meet the forecast. A chi-square test shows that the difference between the two groups is highly significant ($\chi^2 = 32.78$, $p \leq 0.001$). Inferences remain the same in moving to *JMBE* in Panel B. These initial findings suggest that UK firms are more likely to guide analyst forecasts down to

[14] Estimating downward-guided forecasts (*DOWN*) requires two years of lagged data (see equations 2 and 3). As we cannot use pre-FRS 3 observations, to avoid the effect of sample attrition on tests of forecast guidance we replace missing values of *DOWN* by an indicator of negative forecast revisions. This induces measurement error to our proxy to the extent that forecast guidance does not result in negative forecast revisions. In Appendix B we repeat our tests on a restricted sample using the initial estimate of *DOWN*. This leaves the core results unaltered.

**Table 2**

**Panel A: Descriptive statistics for the full sample**

| *Variable* | *N* | *Mean* | *Std. dev* | *25%* | *Median* | *75%* |
|---|---|---|---|---|---|---|
| *MBE* | 5,117 | 60.31% | 48.93% | 0.00% | 100.00% | 100.00% |
| *JMBE* | 3,609 | 65.00% | 47.70% | 0.00% | 100.00% | 100.00% |
| *POSAWCA* | 5,117 | 49.03% | 50.00% | 0.00% | 0.00% | 100.00% |
| *DOWN* | 5,117 | 53.37% | 49.89% | 0.00% | 100.00% | 100.00% |
| *CE* (scaled by sales) | 5,117 | 0.019 | 0.294 | 0.024 | 0.050 | 0.087 |
| $\Delta$*CE* (scales by sales) | 5,117 | 0.003 | 0.188 | –0.007 | 0.006 | 0.020 |
| *UCE* (scaled by sales) | 5,117 | 0.000 | 0.046 | –0.006 | 0.000 | 0.007 |
| $UDCE_{t+1}$ (scaled by sales) | 3,968 | 0.000 | 0.029 | –0.003 | 0.000 | 0.004 |
| *TNRI* (scaled by sales) | 5,117 | 0.021 | 0.103 | 0.000 | 0.000 | 0.008 |
| *NOEI* (scaled by sales) | 5,117 | 0.003 | 0.037 | 0.000 | 0.000 | 0.000 |
| *ONRI* (scaled by sales) | 5,117 | 0.018 | 0.098 | 0.000 | 0.000 | 0.006 |
| *PROFIT* | 5,117 | 90.31% | 29.59% | 100.00% | 100.00% | 100.00% |
| *POS*$\Delta$*EARN* | 5,117 | 61.48% | 48.67% | 0.00% | 100.00% | 100.00% |
| *VREARN* | 5,117 | 3.495 | 2.414 | 1.000 | 3.000 | 5.000 |
| *GROWTH* | 5,117 | 2.843 | 4.250 | 1.130 | 1.843 | 3.080 |
| *LIT* | 5,117 | 18.68% | 38.98% | 0.00% | 0.00% | 0.00% |
| *INDPROD* | 5,117 | 0.009 | 0.025 | 0.000 | 0.007 | 0.019 |
| *SIZE* | 5,117 | 4.501 | 2.868 | 2.000 | 5.000 | 7.000 |

**Table 2** *(continued)*

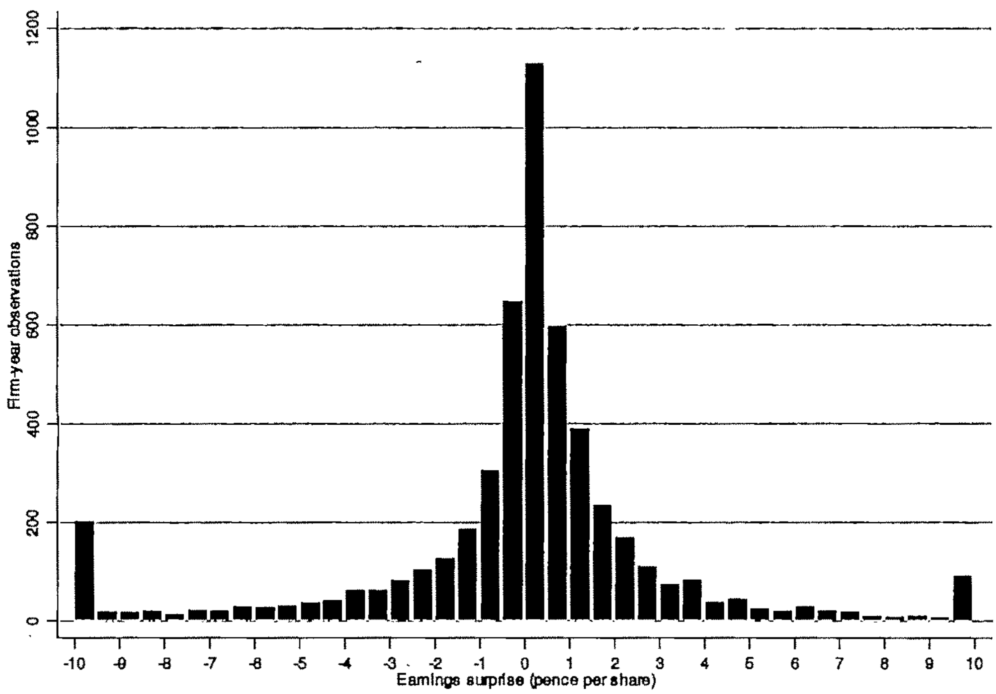
**Panel B: Pearson (above the diagonal) and Spearman (below the diagonal) correlations between key variables**

| | *MBE* | *POSAWCA* | *DOWN* | *CE* (scaled by sales) | *ΔCE* (scaled by sales) | *UCE* (scaled by sales) | $U\Delta CE_{t+1}$ (scaled by sales) | *TNRI* (scaled by sales) | *NOEI* (scaled by sales) | *ONRI* (scaled by sales) |
|---|---|---|---|---|---|---|---|---|---|---|
| *MBE* | | –0.021 | 0.080* | 0.110* | 0.116* | 0.088* | 0.007 | –0.011 | 0.024* | 0.006 |
| *POSAWCA* | –0.021 | | –0.018 | –0.002 | 0.034* | –0.010 | 0.005 | –0.032* | 0.000 | –0.036* |
| *DOWN* | 0.080* | –0.018 | | –0.015 | –0.097* | –0.048* | 0.021 | 0.013 | 0.011 | 0.005 |
| *CE* (scaled by sales) | 0.225* | 0.019 | –0.061* | | 0.367* | 0.177* | –0.041* | –0.010 | –0.034* | 0.014 |
| *ΔCE* (scales by sales) | 0.296* | 0.043* | –0.233* | 0.465* | | 0.297* | –0.023 | –0.020 | –0.041* | 0.012 |
| *UCE* (scaled by sales) | 0.144* | –0.034* | –0.105* | 0.233* | 0.347* | | –0.026 | 0.012 | –0.029* | 0.033* |
| $U\Delta CE_{t+1}$ (scaled by sales) | 0.002 | 0.022 | 0.003 | 0.034* | 0.014 | –0.004 | | –0.001 | 0.004 | –0.011 |
| *TNRI* (scaled by sales) | 0.017 | –0.026* | 0.021 | –0.096* | –0.110* | –0.012 | –0.006 | | 0.592* | 0.642* |
| *NOEI* (scaled by sales) | –0.075* | –0.002 | 0.027* | –0.149* | –0.116* | –0.051* | 0.025 | 0.479* | | –0.205* |
| *ONRI* (scaled by sales) | 0.073* | –0.032* | 0.008 | 0.009 | –0.046* | 0.029* | –0.042* | 0.526* | –0.261* | |

* indicates significance at 10% (two-tailed).

The sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms meeting the sample selection criteria. All observations outside the interval –£0.02 ≤ *SURP* < £0.02 are deleted when *JMBE* is used, resulting in 1,033 firms and 3,609 observations. *SURP* is the earnings surprise measured as the difference between I/B/E/S actual EPS and the latest analyst forecast made prior to the earnings announcement date. Appendix A defines the remaining variables. Calculating unexpected change in core earnings one year ahead reduces the sample to 941 firms and 3,968 observations. The 25% quartile indicates the value of the 25th percentile of the frequency distribution (i.e. the first quarter), while the 75% quartile indicates the value of the 75th percentile of the frequency distribution (i.e. the third quarter).

**Figure 1**
**Frequency distribution of earnings surprises**




The sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms meeting the sample selection criteria. The earnings surprise is the difference between I/B/E/S actual EPS and the latest analyst forecast made prior to the earnings announcement date. We aggregate earnings surprises observations into equally sized intervals (bins). The size of each bin is 0.5p. All earnings surprises below –10p are aggregated to the 20th bin below zero and all earnings surprises above 10p are aggregated to the 20th bin above zero.

meet analyst expectations rather than use income-increasing AWCAs.

*5.2.2. Multivariate analysis*

Model 1 in Table 4 presents results from logistic regressions of *MBE* on indicators of positive AWCAs (*POSAWCA*) and of downward-guided analyst forecasts (*DOWN*), and a series of incentives and controls (equation 5). The Table also reports the marginal effects (economic significance) of the explanatory variables. *POSAWCA* is negative and significant (–0.187, $z = -2.99$). *POSAWCA* is also negative and significant (–0.188, $z = -3.05$) when excluding *DOWN* from the equation. Given the inclusion of *PROFIT* and *POSΔEARN* in this specification, this result suggests that, controlling for the targets of positive earnings levels and changes, firms with income-increasing AWCAs have a lower probability of meeting analyst expectations of about 4%. In a specification excluding *PROFIT* and *POSΔEARN*, the association between *POSAWCA* and *MBE* is insignificant. On the other hand, the coefficient on *DOWN* is positive and highly significant (0.744, $z = 11.07$), consistent with downward-guided forecasts increasing the probability of meeting analyst expectations by approximately 18%. *DOWN* remains positive and significant when excluding *POSAWCA* from the equation (0.743, $z = 11.09$) and in a specification excluding *PROFIT* and *POSΔEARN* (0.297, $z = 5.03$). With the exception of *LIT* and *SIZE* all other variables are significant and in accordance with the predicted signs.[15] Repeating the tests using price volatility (market β) to proxy for risk of shareholder litigation gives an insignificant coefficient on *LIT*. These results are consistent with the less litigious UK business environment. In a specification excluding *PROFIT* and *POSΔEARN*, *SIZE* is positive and significant, consistent with

[15] In a specification that includes $MBE_{t-2}$ and $MBE_{t-3}$ (735 firms, 2,544 observations), while the coefficient on $MBE_{t-1}$ remains positive and significant, the coefficients on additional lags are insignificant. This finding suggests that only last period performance (i.e. achieving analyst expectations in the previous year) affects a firm's ability to meet analyst expectations in the current period.

**Table 3**
**The association between the probability of meeting expectations (*MBE*) and of just meeting expectations (*JMBE*) with earnings management and forecast guidance.**

**Panel A: Contingency tables classifying observations based on: (1) an indicator of meeting analyst forecasts (*MBE*); and (2) indicators of positive abnormal working capital accruals (*POSAWCA*) and of downward-guided analyst forecasts (*DOWN*).**

| | *POSAWCA* = 0 | *POSAWCA* = 1 | *DOWN* = 0 | *DOWN* = 1 |
|---|---|---|---|---|
| *MBE* = 0 | 1,009<br>(38.69%) | 1,022<br>(40.73%) | 1,047<br>(43.88%) | 984<br>(36.03%) |
| *MBE* = 1 | 1,599<br>(61.31%) | 1,487<br>(59.27%) | 1,339<br>(56.12%) | 1,747<br>(63.97%) |
| Total | 2,608 | 2,509 | 2,386 | 2,731 |
| | $\chi^2 = 2.23$<br>$p = 0.135$ | | $\chi^2 = 32.78$<br>$p = \leq 0.001$ | |

**Panel B: Contingency tables classifying observations based on: (1) an indicator of just meeting analyst forecasts (*JMBE*); and (2) indicators of positive abnormal working capital accruals (*POSAWCA*) and of downward-guided analyst forecasts (*DOWN*).**

| | *POSAWCA* = 0 | *POSAWCA* = 1 | *DOWN* = 0 | *DOWN* = 1 |
|---|---|---|---|---|
| *JMBE* = 0 | 622<br>(34.16%) | 641<br>(35.85%) | 639<br>(37.52%) | 624<br>(32.74%) |
| *JMBE* = 1 | 1,199<br>(65.84%) | 1,147<br>(64.15%) | 1,064<br>(62.48%) | 1,282<br>(67.26%) |
| Total | 1,821 | 1,788 | 1,703 | 1,906 |
| | $\chi^2 = 1.15$<br>$p = 0.286$ | | $\chi^2 = 9.05$<br>$p = 0.003$ | |

The sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms that meet the sample selection criteria. Appendix A defines the variables. All observations outside the interval –£0.02 ≤ *SURP* < £0.02 are deleted when *JMBE* is used, resulting in 1,033 firms and 3,609 observations. *SURP* is the earnings surprise measured as the difference between I/B/E/S actual EPS and the latest analyst forecast made prior to the earnings announcement date.

larger firms having stronger incentives to meet analyst expectations.[16]

In Model 2 of Table 4 we repeat the analysis using *JMBE* instead of *MBE* (equation 6). *POSAWCA* is now marginally significant (–0.132, $z = -1.81$), while *DOWN* remains positive and significant at 1% (0.492, $z = 6.22$). Taken together, the results in Table 4 suggest that unlike income-increasing AWCAs, earnings forecast guidance is a common mechanism that UK firms use to avoid negative earnings surprises.

As an additional test of earnings forecast guidance, we follow Bartov et al. (2002) and test the effect of forecast revisions on the sign of earnings surprises. Bartov et al. (2002) argue that in the absence of a forecast revision, the earnings surprise and the quarterly forecast error have the same sign. Accordingly, a negative forecast error that, due to

[16] In addition to litigation risk, value-relevance of earnings and the firm's growth prospects, Matsumoto (2002) examines institutional ownership as a managerial incentive to avoid negative earnings surprises. To the extent institutional investors focus on short-term earnings, and especially on earnings surprises, as they represent a basic metric on which to base trades, managers of firms with higher institutional ownership are more likely to take actions to meet analyst expectations. In line with her prediction, Matsumoto finds that high percentages of institutional ownership increase the probability of meeting or beating analyst expectations. As data on institutional ownership of UK firms are not available for our entire sample, we repeated our core tests adding a control for the percentage of shares held by institutions on a subset of 375 firms (635 observations) that had available ownership structure data and met our sample selection criteria. We are grateful to Dr. Steven Young, Lancaster University, for providing the ownership structure data. On this subset, consistent with Matsumoto (2002), the association between institutional ownership and *MBE* is positive, but only marginally significant (0.747, $z = 1.75$). Future research could assess whether this weak evidence of an institutional impact is due to the effect of different types of institutional holdings in UK firms. While *POSAWCA* is not significant, *DOWN* remains positive and significant (1.056, $z = 5.12$).

**Table 4**
**Logit analysis of the probability of meeting analyst forecasts as a function of indicators of positive abnormal working capital accruals (*POSAWCA*) and of downward-guided analyst forecasts (*DOWN*) and a series of other incentives and controls.**

$$Prob(MBE = 1|X) = F(\beta_0 + \beta_1 POSAWCA_{i,t} + \beta_2 DOWN_{i,t} + \beta_3 PROFIT_{i,t} + \beta_4 POS\Delta EARN_{i,t} + \beta_5 VREARN_{i,t} + \beta_6 GROWTH_{i,t} + \beta_7 LIT_{i,t} + \beta_8 INDPROD_{i,t} + \beta_9 SIZE_{i,t} + \beta_{10} MBE_{i,t-1} + u_{i,t}) \quad (5)$$

$$Prob(JMBE = 1|X) = F(\beta_0 + \beta_1 POSAWCA_{i,t} + \beta_2 DOWN_{i,t} + \beta_3 PROFIT_{i,t} + \beta_4 POS\Delta EARN_{i,t} + \beta_5 VREARN_{i,t} + \beta_6 GROWTH_{i,t} + \beta_7 LIT_{i,t} + \beta_8 INDPROD_{i,t} + \beta_9 SIZE_{i,t} + \beta_{10} MBE_{i,t-1} + u_{i,t}) \quad (6)$$

where $$F(\beta' X) = \frac{e^{\beta' X}}{1 + e^{\beta' X}}$$

| Variable | Predicted sign | *MBE* Coefficient (z-stat) | Marginal effect | *JMBE* Coefficient (z-stat) | Marginal effect |
|---|---|---|---|---|---|
| *Intercept* | | –1.829*** | | –1.336 | |
| | | (–2.08) | | (–1.58) | |
| *POSAWCA* | ? | –0.187*** | –0.044 | –0.132* | –0.030 |
| | | (–2.99) | | (–1.81) | |
| *DOWN* | + | 0.744*** | 0.176 | 0.492*** | 0.111 |
| | | (11.07) | | (6.22) | |
| *PROFIT* | + | 1.293*** | 0.312 | 0.432** | 0.102 |
| | | (9.84) | | (2.29) | |
| *POSΔEARN* | + | 1.253*** | 0.296 | 0.887*** | 0.205 |
| | | (17.94) | | (10.67) | |
| *VREARN* | + | 0.035** | 0.008 | 0.022 | 0.005 |
| | | (2.41) | | (1.32) | |
| *GROWTH* | + | 0.019** | 0.004 | 0.020* | 0.004 |
| | | (2.27) | | (1.97) | |
| *LIT* | + | –0.110 | –0.026 | 0.129 | 0.029 |
| | | (–0.20) | | (0.21) | |
| *INDPROD* | + | 7.339*** | 1.740 | 5.620*** | 1.262 |
| | | (5.61) | | (3.60) | |
| *SIZE* | + | 0.013 | 0.003 | 0.033** | 0.007 |
| | | (1.08) | | (2.35) | |
| $MBE_{t-1}$ | + | 0.320*** | 0.076 | 0.288*** | 0.065 |
| | | (4.98) | | (3.79) | |
| *Industry dummies* | | Yes | | Yes | |
| Log likelihood | | –3,039.14 | | –2,216.47 | |
| Chi-square | | 643.84 | | 220.84 | |
| *p*-value | | <0.001 | | <0.001 | |
| Correctly classified | | 68.88% | | 66.99% | |
| No. of observations: | | | | | |
| Meet | | 3,086 | | 2,346 | |
| Did not meet | | 2,031 | | 1,263 | |
| Total | | 5,117 | | 3,609 | |

*/**/*** indicate significance at 10%/5%/1% (two-tailed). z-statistics in parentheses are based on White standard errors. The marginal effects are computed as $\beta \frac{e^{\beta' X}}{(1+e^{\beta' X})^2}$ where $\beta' X$ is computed at the mean values of X (explanatory variables).

The sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms meeting the sample selection criteria. Appendix A defines the variables. All observations outside the interval –£0.02 ≤ SURP < £0.02 are deleted when *JMBE* is used, resulting in 1,033 firms and 3,609 observations. *SURP* is the earnings surprise measured as the difference between I/B/E/S actual EPS and the latest analyst forecast made prior to the earnings announcement date.

a sufficiently large downward revision, results in a positive earnings surprise is consistent with earnings forecast guidance, whereas a zero or positive forecast error that, due to a sufficiently large upward forecast revision, results in a negative earnings surprise is inconsistent with earnings forecast guidance. Bartov et al. argue that absent management intervention, the proportion of observations where the revision offsets the sign of the earnings surprise should be identical between cases with negative and positive errors. Results of our additional tests show that the frequency of negative forecast errors ending with zero or positive earnings surprises (28%) is substantially higher than the frequency of positive or zero forecast errors ending with negative earnings surprises (7%). This evidence is consistent with earnings forecast guidance to meet analyst expectations, lending further credence to our core findings.

### 5.3. Results on classification shifting to meet analyst expectations

#### 5.3.1. Main results

Table 5 reports the results of regressions of unexpected core earnings on non-recurring items for the entire sample (1,154 firms, 5,117 observations) and three subsets of firms: (a) firms with income increasing non-recurring items (962 firms, 2,803 observations); (b) firms where classifications of non-recurring items allow managers to just hit analyst forecasts (481 firms, 761 observations); and (c) firms where classifications of small non-recurring items allow managers to just hit analyst forecasts (186 firms, 225 observations). Panel A reports the results of equation (9). To the extent UK firms engage in classification shifting, we expect the association between income-increasing non-recurring items and unexpected core earnings to be positive and more profound within the subsets of firms that are increasingly likely to engage in classification shifting to meet analyst expectations. We find no evidence of a significant association between *TNRI* and unexpected core earnings (*UCE*), even focusing on firms where classifications of small total non-recurring items allow them to just hit analyst forecasts. Contrary to the findings of McVay (2006) for US firms, these results suggest that income-increasing total non-recurring items are not systematically related to the variation in unexpected core earnings and thus that classification shifting is not common practice among UK firms.

Some firms might have greater incentives or ability to use classification shifting to meet analyst expectations. Prior research in the UK suggests that larger firms are more likely to engage in classificatory income smoothing to reduce the costs of public visibility (Beattie et al., 1994). Disclosures of alternative earnings metrics are pervasive amongst the 500 largest (based on market capitalisation) UK listed non-financial firms (Choi et al., 2005). Larger firms also experience more frequent structural changes, which often give rise to non-recurring items. We therefore look for evidence of classification shifting on a subset of larger firms. We consider larger firms as those in the highest three quintiles of lagged market capitalisation so that our subset of larger firms resembles the sample of Choi et al. (2005). Within this subset the frequency of zero non-recurring items (8%) is substantially lower than that for the remaining firms in the sample (16%), consistent with larger firms disclosing non-recurring items more frequently. Repeating equation (9) for larger firms where classifications of small total non-recurring items allow them to just hit analyst forecasts (113 firms, 137 observations), *TNRI* remains insignificant.

Panel B of Table 5 reports the results of equation (11) on the separate components of *TNRI* (*NOEI* and *ONRI*). As other non-recurring items offer a more flexible device for classification shifting to meet analyst expectations than non-operating exceptional items, we expect the coefficient on *ONRI* to be positive. On the entire sample, firms with income-increasing other non-recurring items (962 firms, 2,927 observations) and firms where classifications of other non-recurring items allow managers to just hit analyst forecasts (489 firms, 761 observations), we find no evidence of a significant association between *UCE* and income-increasing *ONRI*. However, for firms where classifications of small other non-recurring items allow them to just hit the forecast (201 firms, 244 observations), *ONRI* is positive and significant (0.597, $t$=2.88). This is due to larger firms within this sub-sample, as for larger firms that would have just missed the target without classifications of small non-recurring items (121 firms, 148 observations) the coefficient on *ONRI* is positive and significant (0.868, $t$=2.06). For these firms, a one standard deviation increase in other non-recurring items results in an increase in unexpected core earnings (scaled by sales) of 26 basis points (0.868, multiplied by 0.003, the standard deviation of other non-recurring items for this sample). Not surprisingly, as the sample is narrowed to focus on firms that have greater opportunity to engage in classification shifting, the adjusted $R^2$ increases from to 0.2% to 1.3%. As expected, we find no evidence of a positive association between *UCE* and income-increasing *NOEI* for any subset of firms.

Summarising, the results in Table 5 provide evidence of a positive association between unexpected core earnings and small other non-recurring items for larger firms that met the analyst forecast by up to two pence per share, when otherwise they would have just missed the target. This evidence is

**Table 5**
**Unexpected core earnings and income-increasing non-recurring items**

**Panel A: Regression of unexpected core earnings on total non-recurring items as a percentage of sales, for the entire sample and subsets of firms that are increasingly likely to use classification shifting to meet analyst expectations.**

$$UCE_{i,t} = \delta_0 + \delta_1 TNRI_{i,t} + u_{i,t}$$

| Variable | Predicted sign | *Entire sample* | *Income-increasing TNRI* | *JUSTMET with TNRI* | *JUSTMET with Small TNRI* | *JUSTMET with Small TNRI (larger firms)* |
|---|---|---|---|---|---|---|
| *Intercept* | | 0.000 | 0.001 | 0.000 | –0.003 | –0.003 |
| | | (0.12) | (0.78) | (–0.12) | (–0.99) | (–0.66) |
| *TNRI* | + | 0.014 | 0.012 | 0.014 | 0.506 | 0.624 |
| | | (1.08) | (0.89) | (0.39) | (0.80) | (0.98) |
| *Adjusted $R^2$* | | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| Number of firms | | 1,154 | 962 | 481 | 186 | 113 |
| Number of observations | | 5,117 | 2,803 | 761 | 225 | 137 |

**Table 5**
**Unexpected core earnings and income-increasing non-recurring items** *(continued)*

**Panel B: Regression of unexpected core earnings on non-operating exceptional and other non-recurring items as a percentage of sales for the entire sample and subsets of firms that are increasingly likely to use classification shifting to meet analyst expectations.**

$$UCE_{i,t} = \delta_0 + \delta_{1a}NOEI_{i,t} + \delta_{2b}ONRI_{i,t} + u_{i,t}$$

| Variable | Predicted sign | *Entire sample* | *Income-increasing ONRI* | *JUSTMET with ONRI* | *JUSTMET with small ONRI* | *JUSTMET with small ONRI (larger firms)* |
|---|---|---|---|---|---|---|
| *Intercept* | | 0.000 | 0.002** | 0.001 | –0.001 | –0.001 |
| | | (0.27) | (2.47) | (0.47) | (–0.59) | (–0.22) |
| *NOEI* | + | –0.087** | –0.021 | 0.014 | –0.008 | –0.026 |
| | | (–2.12) | (–0.69) | (0.35) | (–0.17) | (–0.64) |
| *ONRI* | + | 0.030 | 0.020 | 0.023 | 0.597*** | 0.868** |
| | | (1.60) | (1.10) | (0.58) | (2.88) | (2.06) |
| *Adjusted* $R^2$ | | 0.009 | 0.002 | 0.002 | 0.006 | 0.013 |
| Number of firms | | 1,154 | 962 | 489 | 201 | 121 |
| Number of observations | | 5,117 | 2,927 | 761 | 244 | 148 |

*/**/*** indicate significance at 10%/5%/1% (two-tailed). *t*-statistics in parentheses are based on robust standard errors clustered by year.
The entire sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms meeting the sample selection criteria. Appendix A defines the variables. *JUSTMET with TNRI* includes observations where the earnings surprise (*SURP*) is in the interval $-£0.02 \le SURP < £0.02$, total non-recurring items are income-increasing ($TNRI>0$) and $SURP - TNRI < 0$. *JUSTMET with Small TNRI* includes observations where the *SURP* is in the interval $-£0.02 \le SURP < £0.02$, $TNRI>0$, *TNRI* is less than median *TNRI* for firms with positive *TNRI* and $SURP - TNRI < 0$. *JUSTMET with ONRI* includes observations where *SURP* is in the interval $-£0.02 \le SURP < £0.02$, other non-recurring items are income-increasing ($ONRI>0$) and $SURP - ONRI < 0$. *JUSTMET with Small ONRI* includes observations where the *SURP* is in the interval $-£0.02 \le SURP < £0.02$, $ONRI>0$, *ONRI* is less than median *ONRI* for firms with positive *ONRI* and $SURP - ONRI < 0$. Larger firms are those in the highest three quintiles of lagged market capitalisation.

consistent with a subset of larger firms engaging in classification shifting of small core expenses to operating exceptional or other non-recurring items to just meet analyst expectations. On the other hand, a positive association between *UCE* and *ONRI* may also result from immediate efficiency gains due to rationalising operations. To distinguish between the two scenarios, we next examine the association between non-recurring items and future profitability.

Table 6 reports results from regressions of future changes in unexpected core earnings (*UΔCE*) on non-recurring items for the entire sample and the three subsets of firms that are increasingly likely to engage in classification shifting to meet analyst expectations. To the extent UK firms engage in classification shifting to meet analyst expectations, we expect the association between non-recurring items and changes in unexpected core earnings one year ahead to be negative and more profound for the sub-samples. Panel A reports the results of equation (10). For firms where classifications of small total non-recurring items allow them to just hit the forecast (155 firms, 178 observations), *TNRI* is negative and significant (–0.553, $t$=–2.28). The negative association pertains to larger firms within this subset (94 firms, 113 observations), as for these firms *TNRI* is significantly negative and higher in magnitude (–1.160, $t$ = –3.02).[17] This result is more likely due to the effect of small income-increasing other non-recurring items, as for larger firms that just hit the analyst forecast with classifications of these items (Panel B, 102 firms, 120 observations) *ONRI* is negative (–1.060, $t$=–1.68). Combined with the evidence in Table 5, this finding suggests that for larger firms that just meet analyst forecasts by shifting small core expenses to other non-recurring items, a one standard deviation increase in other non-recurring items is expected to increase unexpected core earnings by 26 basis points and have a subsequent incremental reversal of about 22 basis points (–1.060 multiplied by 0.002) one year ahead. Almost 85% (22/26) of the increase in unexpected core earnings reverses in the subsequent period, as these expenses recur. Taken together, the results in Tables 5 and 6 provide evidence that a subset of larger firms, roughly 10% of sample firms (121/1154), engage in classification shifting of small other non-recurring items to just meet analyst expectations.[18]

Despite the increase in adjusted $R^2$ for firms that are more likely to engage in classification shifting, the explanatory power for all regressions in Tables 5 and 6 is low (0.01%–5.1%). In similar regressions, McVay (2006: 518–521) also reports low adjusted $R^2s$ (0.01%–8.5%), despite her larger dataset and the stronger evidence of an association between non-recurring items and unexpected core profits for US firms. The low explanatory power of these tests is due to the two-stage process that we follow, after McVay (2006). For example, in the first stage given by equation (7) we model expected core earnings by industry–year; in the second stage (equation 9) we run a simple regression to test the association between unexpected core earnings and non-recurring items. This two-stage design yields a low explanatory power in the second stage, but our overall ability to explain variation in core earnings is determined by both the first and second stages. The low explanatory power of the second stage is also due to the positive (negative) association between unexpected core profits and non-recurring items pertaining to a small subset of larger firms.

### 5.3.2. *Additional analysis*

For a complete assessment of the implications of non-recurring items for firms that are most likely to be engaging in classification shifting to meet analyst expectations, we test the association between non-recurring items and future operating cash flows:

$$FutureCFO_{i,t} = \lambda_0 + \lambda_1 CE_{i,t} + \lambda_2 NOEI_{i,t} + \lambda_3 ONRI_{i,t} + \lambda_4 ACCRUALS_{i,t} + \lambda_5 GROWTH_{i,t} + u_{i,t} \quad (13)$$

where *FutureCFO* is operating cash flows summed over three years following year $t$ and scaled by sales. If non-recurring items predict future cash flows, the precise timing of this effect in terms of the window for measuring future cash flows is uncertain. If managers shift core expenses to non-recurring items to meet expectations, they can use strategic disclosures to show higher core earnings in the short term, but they cannot avoid negative implications in the long term. As a result, we sum operating cash flows over three years to ensure we capture a substantial portion of the predictive power of non-recurring items. We split total non-recurring items into *NOEI* and *ONRI* to examine the separate effects of the two components. Similar to Doyle et al. (2003), we control

[17] As tests of future unexpected changes in core earnings require one-year-ahead data, reducing the sample to 941 firms and 3,968 observations, these results are not directly comparable with results in Table 5, which are calculated on the original sample of 1,154 firms and 5,117 observations. Repeating the analysis in Table 5 on the subset of 3,968 observations leaves the core findings unaltered.

[18] An alternative approach to clustering by year to control for cross sectional dependence is to use Fama and MacBeth (1973) across year $t$-statistics. However, as we run equations (9)–(12) on gradually narrower subsets of firms (the last subset includes 16 observations on average per year) the Fama–MacBeth statistics may bias results against finding an association between income-increasing non-recurring items and unexpected core profits. Indeed this alternative estimation yields weaker evidence of classification shifting within the subset of larger firms that just hit analyst forecast using small income-increasing other non-recurring items.

**Table 6**
**Unexpected changes in core earnings and income-increasing, non-recurring items**

**Panel A: Regression of future unexpected changes in core earnings on total non-recurring items as a percentage of sales for the entire sample and subsets of firms that are increasingly likely to use classification shifting to meet analyst expectations.**

$$U\Delta CE_{i,t+1} = \delta_0 + \delta_1 TNRI_{i,t} + u_{i,t}$$

| Variable | Predicted sign | *Entire sample* | *Income-increasing TNRI* | *JUSTMET with TNRI* | *JUSTMET with Small TNRI* | *JUSTMET with Small TNRI (larger firms)* |
|---|---|---|---|---|---|---|
| *Intercept* | | 0.000<br>(0.05) | –0.001<br>(–0.56) | 0.001<br>(0.46) | 0.002<br>(1.22) | 0.003<br>(1.28) |
| *TNRI* | – | 0.000<br>(–0.01) | 0.002<br>(0.14) | 0.016<br>(0.55) | –0.553**<br>(–2.28) | –1.160***<br>(–3.02) |
| *Adjusted $R^2$* | | 0.000 | 0.000 | 0.003 | 0.008 | 0.051 |
| Number of firms | | 941 | 796 | 401 | 155 | 94 |
| Number of observations | | 3,968 | 2,134 | 609 | 178 | 113 |

**Table 6**
**Unexpected changes in core earnings and income-increasing, non-recurring items** *(continued)*

**Panel B: Regression of future unexpected changes in core earnings on non-operating exceptional and other non-recurring items as a percentage of sales for the entire sample and subsets of firms that are increasingly likely to use classification shifting to meet analyst expectations.**

$$U\Delta CE_{i,t+1} = \delta_0 + \delta_{1a}NOEI_{i,t} + \delta_{2b}ONRI_{i,t} + u_{i,t}$$

| Variable | Predicted sign | *Entire sample* | *Income-increasing ONRI* | *JUSTMET with ONRI* | *JUSTMET with Small ONRI* | *JUSTMET with Small ONRI (larger firms)* |
|---|---|---|---|---|---|---|
| *Intercept* | | 0.000 | –0.001 | 0.001 | 0.001 | 0.002 |
| | | (0.29) | (–1.14) | (0.68) | (0.80) | (0.84) |
| *NOEI* | – | 0.028 | –0.004 | –0.002 | –0.054 | –0.064 |
| | | (1.53) | (–0.11) | (–0.12) | (–0.77) | (–1.21) |
| *ONRI* | – | –0.012 | –0.007 | –0.006 | –0.124 | –1.060* |
| | | (–1.14) | (–0.64) | (–0.38) | (–0.27) | (–1.68) |
| *Adjusted $R^2$* | | 0.002 | 0.001 | 0.001 | 0.002 | 0.019 |
| Number of firms | | 941 | 794 | 402 | 161 | 102 |
| Number of observations | | 3,968 | 2,262 | 600 | 188 | 120 |

*/**/*** indicate significance at 10%/5%/1% (two-tailed). t-statistics in parentheses are based on robust standard errors clustered by year.
The original sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms meeting the sample selection criteria. Calculating unexpected change in core earnings requires one year ahead of data reducing the sample to 941 firms and 3,968 observations. Appendix A defines the variables. *JUSTMET with TNRI* includes observations where the earnings surprise (*SURP*) is in the interval –£0.02 ≤ *SURP* < £0.02, total non-recurring items are income-increasing (*TNRI*>0) and *SURP* – *TNRI* < 0. *JUSTMET with Small TNRI* includes observations where the *SURP* is in the interval –£0.02 ≤ SURP < £0.02, *TNRI*>0, *TNRI* is less than median *TNRI* for firms with positive *TNRI* and *SURP* – *TNRI* < 0. *JUSTMET with ONRI* includes observations where *SURP* is in the interval –£0.02 ≤ *SURP* < £0.02, other non-recurring items are income-increasing (*ONRI*>0) and *SURP* – *ONRI* < 0. *JUSTMET with Small ONRI* includes observations where the *SURP* is in the interval –£0.02 ≤ *SURP* <£0.02, *ONRI*>0, *ONRI* is less than median *ONRI* for firms with positive *ONRI* and *SURP* – *ONRI* < 0. Larger firms are those in the highest three quintiles of lagged market capitalisation.

for *GROWTH* and *ACCRUALS*. Higher-growth firms have higher changes in working capital, which may affect future cash flows. Also higher-growth firms tend to experience structural events, which are the major source of non-operating exceptional items. We add accruals as prior research shows that they predict future cash flows (Dechow, 1994; Sloan, 1996; Barth et al., 2001). If an income-increasing non-recurring item is booked as an accrual (e.g. asset impairment, provision for reorganisation costs) in the current period and turns into a cash outflow in a future period, then accruals reversal could yield a negative association between non-recurring items and future operating cash flows.

Similar to equations (11) and (12), we estimate equation (13) for the entire sample and four subsets of firms: (a) firms with income-increasing other non-recurring items; (b) firms where classifications of other non-recurring items allow managers to just hit analyst forecasts; (c) firms where classifications of small other non-recurring items allow managers to just hit analyst forecasts; and (d) larger firms from subset three (the highest three quintiles of lagged market capitalisation). If, consistent with evidence in Tables 5 and 6, managers of larger firms engage in classification shifting of small core expenses to other non-recurring items to just meet analyst expectations, then we expect the coefficient on *ONRI*, $\lambda_3$, to be negative for the fourth subset of firms, capturing the effect of small recurring expenses on these firms' operating cash flows in subsequent accounting periods.

For equation (13), aggregating operating cash flows three years ahead reduces the sample to 863 firms and 3,427 observations. We mitigate sample attrition by collecting data on operating cash flows for 2003 and 2004. Because the dependent variable in equation (13) aggregates operating cash flows over three years, it involves overlapping observations. To address serial correlation in error terms further to cross-sectional dependence, we estimate standard errors clustered by firm and year.[19] Panel A of Table 7 reports the regression results. The results show that other non-recurring items are not associated with future cash flows for the entire sample, the subset of firms with income-increasing *ONRI* and the subset of firms just hitting analyst forecasts with income-increasing *ONRI*. However, for firms that just hit analyst forecasts using classifications of small income-increasing *ONRI* (subset 3, 140 firms, 161 observations) and larger firms within this subset (subset 4, 86 firms, 98 observations), other non-recurring items are recurring and consume cash in the future. The coefficients on *ONRI* are negative and significant for both subset 3 (–16.739, $t$=–2.69) and subset 4 (–13.417, $t$=–2.76). As expected, future operating cash flows generally decline with *ACCRUALS*. Since *CE*, *NOEI* and *ONRI* capture net income, the accrual result shows that given net income, a relatively high level of accruals predicts lower future cash flows, consistent with the results from prior research (Dechow, 1994; Sloan, 1996; Barth et al., 2001). More important, *ACCRUALS* control for the mechanical relation between expenses that accrue in the current period and turn into cash in the future periods when they reverse. Therefore, the coefficients on *ONRI* capture the incremental cash outflows in the future beyond the reversing of accrued expenses. Inspecting the magnitude of the coefficients, *ONRI* is significantly higher in absolute value than *CE* for the two subsets of firms (3,202 and 3.742 respectively). This is most likely due to the small size of other non-recurring items within the two subsets of firms, rather than to a higher degree of permanence of these items compared to core earnings. Based on standardised regression coefficients the degree of permanence of core earnings is substantially higher than that of *ONRI* for both subsets of firms. A one standard deviation increase in *CE* for the third (fourth) subset of firms increases future operating cash flows by 56% (63%), while a one standard deviation increase in *ONRI* for the third (fourth) subset of firms decreases future operating cash flows by 11% (10%).

Overall, results in Table 7 show that even though, on average, non-recurring items do not convey information about future operating performance, for firms that hit the analyst target using classifications of small other non-recurring items, these items predict future cash outflows. These results reinforce our evidence of larger firms engaging in classification shifting of small core expenses to operating exceptional or other non-recurring items to just meet analyst expectations.

[19] Petersen (2007) suggests that when a model suffers from both serial correlation in the disturbance terms and cross-sectional dependence, standard errors clustered on two dimensions (time and firm) are unbiased and produce correctly sized intervals whether the firm effect is permanent or temporary. We make similar inferences when we estimate the equations using the Prais-Winsten estimator (feasible generalised least squares), which corrects for cross-sectional correlation and autocorrelated and heteroscedastic residuals. An alternative approach to alleviating cross sectionally and serially correlated error terms is to use Fama and MacBeth (1973) across year $t$-statistics and multiply the traditional standard error by the Newey–West adjustment factor similar to Doyle et al. (2003). However, as we run equation (13) on gradually narrower subsets of firms (the last subset includes 12 observations on average per year) the Fama–MacBeth statistics may bias results against finding an association between income-increasing non-recurring items and future cash flows. Indeed this alternative estimation yields weaker evidence of a negative association between other non-recurring items and future cash flows for firms that just hit analyst forecast using small income-increasing other non-recurring items.

**Table 7**
**Regressions of future operating cash flows on current core earnings, non-recurring items and a set of control variables**

$$FutureCFO_{i,t} = \lambda_0 + \lambda_1 CE_{i,t} + \lambda_2 NOEI_{i,t} + \lambda_3 ONRI_{i,t} + \lambda_4 ACCRUALS_{i,t} + \lambda_5 GROWTH_{i,t} + u_{i,t}$$

| Variable | *Entire sample* | *Income-increasing ONRI* | *JUSTMET with ONRI* | *JUSTMET with Small ONRI* | *JUSTMET with Small ONRI (larger firms)* |
|---|---|---|---|---|---|
| *Intercept* | 0.215*** | 0.200*** | 0.107*** | 0.110*** | 0.066*** |
| | (8.74) | (8.40) | (3.40) | (2.74) | (2.75) |
| *CE* | 1.628*** | 1.700*** | 2.328*** | 3.202*** | 3.742*** |
| | (12.08) | (17.97) | (4.64) | (7.01) | (12.34) |
| *NOEI* | –0.199 | –0.694 | –0.806** | –0.440 | –0.233 |
| | (–0.65) | (–2.96) | (–2.46) | (–0.58) | (–0.43) |
| *ONRI* | –0.253 | –0.128 | –0.002 | –16.739*** | –13.417*** |
| | (–0.93) | (–0.45) | (–0.01) | (–2.69) | (–2.76) |
| *ACCRUALS* | –1.502*** | –1.388*** | –2.490*** | –1.835*** | –2.250*** |
| | (–5.16) | (–4.45) | (–4.46) | (–13.77) | (–6.09) |
| *GROWTH* | –0.003 | 0.001 | –0.001 | –0.003 | –0.004 |
| | (–1.51) | (0.42) | (–0.33) | (–0.78) | (–0.97) |
| *Adjusted* $R^2$ | 0.5508 | 0.6406 | 0.6638 | 0.6312 | 0.7668 |
| Number of firms | 863 | 703 | 328 | 140 | 86 |
| Number of observations | 3,427 | 1,934 | 495 | 161 | 98 |

*/**/*** indicate significance at 10%/5%/1% (two-tailed). *t*-statistics in parentheses are based on robust standard errors clustered by year and firm to control for cross-sectional dependence and heteroskedastic and autocorrelated residuals.

The original sample consists of 5,117 observations during the period 1994–2002 for 1,154 UK firms meeting the sample selection criteria. Appendix A defines the variables. Calculating *FutureCFO* requires three years ahead of data, reducing the sample to 863 firms and 3,427 observations. *JUSTMET with ONRI* includes observations where *SURP* is in the interval $-£0.02 \leq SURP < £0.02$, other non-recurring items are income-increasing ($ONRI > 0$) and $SURP - ONRI < 0$. *JUSTMET with small ONRI* includes observations where the *SURP* is in the interval $-£0.02 \leq SURP < £0.02$, $ONRI > 0$, *ONRI* is less than median *ONRI* for firms with positive *ONRI* and $SURP - ONRI < 0$. Larger firms are those in the highest three quintiles of lagged market capitalisation.

## 6. Conclusion

This paper sheds light on the mechanisms UK firms use to meet analyst expectations. We examine earnings forecast guidance and two earnings management mechanisms. The results of our investigation over the period 1994–2002 show that downward-guided forecasts increase the probability of meeting analyst expectations. Furthermore, within a small subset of larger firms (based on market capitalisation), we find evidence consistent with classification shifting of small other non-recurring items to hit analyst earnings forecasts. In particular, for larger firms that would have just missed the target without classifications of other non-recurring items, we find that other non-recurring items are associated with an abnormal rise in core profits in the current period, an abnormal decline in core earnings in the subsequent period and operating cash outflows three years ahead. This evidence of classification shifting should be treated with caution as it pertains only to a narrow subset of firms, where the explanatory power of non-recurring items for excess core earnings is low and is weaker in some robustness tests. With regard to the second earnings management mechanism, we find no evidence of a positive association between income-increasing AWCAs and the probability of meeting analyst expectations. Overall, our results suggest that UK firms are more likely to engage in earnings forecast guidance or, for a subset of larger firms, in classification shifting than in accruals management to meet analyst expectations. Our results corroborate the survey evidence of Choi et al. (2006) that UK managers are more likely to engage in earnings forecast guidance to meet earnings benchmarks than to bear the costs of deploying income-increasing discretionary accounting choices.

Our results have important implications for both investors and accounting standard setters. Investors view a zero or positive earnings surprise as evidence of a well-managed firm, able to both predict and deliver future earnings. Graham et al. (2005: 5) remark that, 'The severe stock market reactions to small EPS misses can be explained as evidence that the market believes that most firms can "find the money" to hit earnings targets. Not being able to find one to two cents to hit the target might be interpreted as evidence of hidden problems at the firm.' The market therefore appears to expect a certain degree of earnings or expectations management to hit the target. To the extent UK firms guide analyst forecast down or engage in classification shifting to make the target attainable, the question that arises is whether investors take this into consideration by discounting the market reward (if any) to meeting analyst expectations.

With regard to accounting standard setters, our results suggest that under an FRS 3 regime, UK firms are unlikely to engage in earnings management through accruals. This is in line with the original intention of the UK regulatory bodies in the early 1990s to reduce opaque practices of earnings manipulation. Moreover, while prior research suggests that misclassifications of extraordinary items were pervasive before the introduction of FRS 3 (Smith, 1992), our results show that the practice of classification shifting through non-recurring items was not common post-FRS 3. FRS 3's rigorous transparency requirements, especially with regard to large visible transitory items, contributed to this effect, as they helped users to ascertain with greater confidence the nature of the items firms classify as non-recurring, increasing the cost of using classification shifting to manage earnings. While classification shifting was not widespread post-FRS 3, we find evidence consistent with the practice within a subset of larger UK firms. This evidence suggests that despite the transparency restraints of FRS 3, there was still scope for shifting relatively small recurring expenses to operating exceptional and other non-recurring items. Future research should assess whether investors can detect classification shifting attempts and impound their full implications for stock prices. A further interesting avenue for future research would be to investigate the effect of large audit firms on the use of classification shifting to achieve analyst expectations. This avenue could extend to earnings forecast guidance. A factor that needs to be considered is the extent to which these mechanisms raise auditor suspicion, especially in view of their subtle nature and the role of auditing relative to other corporate governance mechanisms (e.g. institutional ownership, managerial ownership, the monitoring action of non-executive directors, the board of directors and the audit committee).

**Appendix A**
**Definition of variables**

| *Variable* | *Definition* |
|---|---|
| *ACCRUALS* | Accruals scaled by sales. Accruals is the difference between adjusted ordinary earnings (DS210) and operating cash flows from the cash flows statement (DS1015). |
| *AFO* | Latest forecast for the year made prior to the earnings announcement date. |
| *ATO* | Asset turnover defined as total sales over average net operating assets $[(NOA_t+NOA_{t-1})/2]$. *NOA* (net operating assets) is the difference between operating assets and operating liabilities. Operating assets is total assets (DS392) minus cash and cash equivalent (DS375). Operating liabilities is total assets (DS392) less total debt (DS1301), total equity (DS307) and minority interest (DS176). |
| *CE* | Core earnings calculated as I/B/E/S actual EPS multiplied by the weighted average number of shares (both unadjusted for splits to derive the historical figures) and scaled by total sales (DS104). |
| *CRET* | Excess monthly return cumulated from the month following the year *t*-1 earnings announcement to the month of the year *t* earnings announcement. Excess return is firm return less the market return using the FTSE All Shares Index. Returns are collected from the LSPD 2002. |
| *DOWN* | Equals 1 if *UEF* is negative, 0 otherwise. We estimate *EF* based on prior year earnings and cumulative returns during the year. Because the estimation process requires two years of lagged data, the initial estimate of *DOWN* has many missing observations during our sample period. We replace missing values with an indicator of negative forecast revisions. Forecast revision is the difference between the latest forecast (*AFO*) and the earliest forecast for the year (following the announcement of prior year's earnings). |
| *EF* | Expected latest forecast (*AFO*). |
| *EPS* | I/B/E/S reported actual EPS. |
| *FutureCFO* | Operating cash flow (DS1015) summed over three years following the current accounting period and scaled by sales. |
| *GROWTH* | The market value of outstanding shares at the end of the year (DSHMV) divided by book value of common equity at the end of the year (DS307), similar to Skinner and Sloan (2002). |
| *INDPROD* | Average annual growth in industrial production (using UKCKYWA-UK industrial production index: all production industries) adjusted for inflation (using UKCONPRCF-UK RPI index). |
| *JMBE* | Equals 1 if the earnings surprise (*SURP*) is in the interval $0 \leq SURP <$ £0.02, 0 if the earnings surprise is in the interval $-$£0.02 $\leq SURP < 0$. |
| *JUSTMET* | Equals 1 if the earnings surprise is from £0.00 to £0.02 per share, non-recurring items are income-increasing and the earnings surprise minus total non-recurring items per share is negative, 0 otherwise. |
| *LARGER* | Equals 1 for firms in the highest three quintiles of lagged market capitalisation each year, 0 otherwise. |
| *LIT* | Equals 1 if the firm belongs to a high risk industry (biotechnology, computers, electronics and retail), 0 otherwise (see Matsumoto, 2002). High risk industries are Datastream Level 6 BIOTC, CMPSV, INTNT, SOFTW, ELETR, DSCST, ERETL, HARDL, MULTI, SOFTG. |
| *MBE* | Equals 1 if the earnings surprise (*SURP*) is zero or positive, 0 otherwise. *SURP* is the difference between I/B/E/S actual EPS and the latest forecast for the year made prior to the earnings announcement date (*AFO*). In line with Bartov et al. (2002), we choose the latest forecast to precede the earnings release date by at least three days to ensure knowledge of the actual earnings figure does not contaminate the forecast. |

**Appendix A**
**Definition of variables** *(continued)*

| *Variable* | *Definition* |
|---|---|
| *NEG_ΔSALES* | Equals 1 if *ΔSALES* is negative, 0 otherwise. |
| *NOEI* | Non-operating exceptional items adjusted for tax and minority interest (DS1083 – DS1094 – DS1097) and scaled by sales. DS1083 is total non-operating exceptional items and includes profits or losses on the sale or termination of operations, costs of fundamental reorganisations or restructuring and profits or losses on the sale of fixed assets; DS1094 is tax on non-operating exceptional items; and DS1097 is the minority interest on non-operating exceptional items. Datastream records exceptional and extraordinary items as negative when they are costs or losses and positive when they are revenues or gains. To capture the income-increasing (decreasing) effect of removing negative (positive) items, we multiply *NOEI* by –1. |
| *ONRI* | Other non-recurring items calculated as the difference between *TNRI* and *NOEI*. |
| *POSAWCA* | Equals 1 if AWCA is positive, 0 otherwise. Abnormal working capital accruals is estimated using the cross-sectional modified Jones model including lagged return on assets (Kothari, Leone and Wasley, 2005), deleting industry–year combinations with less than six observations. We calculate working capital accruals (WCAs) directly from the cash flow statement as change in debtors (DS448), plus change in inventory (DS444), minus change in creditors (DS417) and minus other changes in working capital (DS1012). By deriving WCA from the cash flow statement, we avoid the potential measurement error in accruals derived from the balance sheet when non-operating events such as mergers, acquisitions and divestitures occur (Collins and Hribar, 2002). |
| *POSΔEARN* | Equals 1 if annual change in I/B/E/S actual EPS is positive, 0 otherwise. |
| *PROFIT* | Equals 1 if I/B/E/S actual EPS is positive in the current accounting period, 0 otherwise. |
| *ΔREC* | Change in accounts receivable (DS448). |
| *ΔREV* | Change in revenue (DS104). |
| *ROA* | Earnings before interest, tax, depreciation and amortisation (DS1502) over total assets (DS392). |
| *ΔSALES* | Change in sales (DS104). |
| *SIZE* | Decile portfolios formed each year by sorting observations into 10 groups based on lagged market value of equity (0 is the lowest, 9 the highest decile). |
| *Small ONRI* | Other non-recurring items (*ONRI*) are less than median *ONRI* for firms with positive *ONRI*. |
| *Small TNRI* | Total non-recurring items (*TNRI*) are less than median *TNRI* for firms with positive *TNRI*. |
| *TNRI* | Total non-recurring items calculated as the difference between *CE* and net income (DS1087) scaled by sales. |
| *UΔCE* | Unexpected change in core earnings scaled by sales. |
| *UCE* | Unexpected core earnings scaled by sales. |
| *UEF* | Unexpected latest analyst forecast (*AF0 – EF*). |
| *VREARN* | Decile portfolios formed each year by sorting $R^2$s from industry (Datastream Level 6) specific regressions of excess returns (cumulated from the month following the year $t$–1 earnings announcement to the month of the year $t$ earnings announcement) on annual change in I/B/E/S actual EPS. Excess returns are firm returns less market returns using the FTSE All Shares Index. Returns are from the LSPD. We assign the value of 0 to firms in the smallest decile through to 9 for firms in the largest decile. |
| *WCA* | Working capital accruals scaled by sales. |

**Appendix B**
**Results on earnings forecast guidance using the initial estimate of *DOWN***

**Panel A: Contingency tables classifying observations based on (1) an indicator of meeting analyst forecasts (*MBE*) or of just meeting analyst forecasts (*JMBE*); and (2) an indicator of downward-guided analyst forecasts (*DOWN*).**

| | *DOWN* = 0 | *DOWN* = 1 | | *DOWN* = 0 | *DOWN* = 1 |
|---|---|---|---|---|---|
| MBE = 0 | 883 (45.61%) | 702 (34.53%) | JMBE = 0 | 530 (38.91%) | 440 (31.29%) |
| MBE = 1 | 1,053 (54.39%) | 1,331 (65.47%) | JMBE = 1 | 832 (61.09%) | 966 (68.71%) |
| Total | 1,936 | 2,033 | | 1,362 | 1,406 |
| | $\chi^2 = 50.75$ $p = <0.001$ | | | $\chi^2 = 17.64$ $p = <0.001$ | |

**Panel B: Logit analysis of the probability of meeting analyst forecasts (*MBE* = 1, *JMBE* = 1) as a function of indicators of positive abnormal working capital accruals (*POSAWCA*) and of downward-guided analyst forecasts (*DOWN*) and a series of incentives and controls.**

| Variable | Predicted sign | *MBE* Coefficient (z-stat) | Marginal effect | *JMBE* Coefficient (z-stat) | Marginal effect |
|---|---|---|---|---|---|
| *Intercept* | | −2.178*** (−9.24) | | −0.930*** (−3.01) | |
| *POSAWCA* | ? | −0.204*** (−2.86) | −0.049 | −0.118 (−1.43) | −0.027 |
| *DOWN* | + | 0.856*** (11.10) | 0.201 | 0.610*** (6.73) | 0.136 |
| *PROFIT* | + | 1.206*** (7.77) | 0.293 | 0.262 (1.13) | 0.061 |
| *POSΔEARN* | + | 1.365*** (17.12) | 0.321 | 0.997*** (10.55) | 0.230 |
| *VREARN* | + | 0.018 (1.06) | 0.004 | 0.004 (0.38) | 0.001 |
| *GROWTH* | + | 0.003 (0.36) | 0.001 | 0.005 (0.48) | 0.001 |
| *LIT* | + | −0.143 (−0.14) | −0.034 | 0.627 (0.60) | 0.130 |
| *INDPROD* | + | 6.647*** (3.75) | 1.577 | 3.806* (1.70) | 0.854 |
| *SIZE* | + | 0.007 (0.57) | 0.002 | 0.030* (1.85) | 0.007 |
| $MBE_{t-1}$ | + | 0.346*** (4.65) | 0.083 | 0.318*** (3.60) | 0.072 |
| *Industry dummies* | | Yes | | Yes | |

**Appendix B**
**Results on earnings forecast guidance using the initial estimate of *DOWN* (*continued*)**

**Panel B continued: diagnostic statistics**

| | *MBE* | *JMBE* |
|---|---|---|
| Log likelihood | –2,343.09 | –1,688.20 |
| Chi-square | 540.88 | 191.57 |
| *p*-value | <0.001 | <0.001 |
| Correctly classified | 69.44% | 67.31% |
| No. of observations: | | |
| Meet | 2,384 | 1,798 |
| Did not meet | 1,585 | 970 |
| Total | 3,969 | 2,768 |

*/**/*** indicate significance at 10%/5%/1% (two-tailed). z-statistics in parentheses are based on White standard errors.
The sample consists of 3,969 observations during the period 1995–2002 for 937 UK firms meeting the sample selection criteria and having available data for *DOWN*. *DOWN* equals 1 if the unexpected part of the latest forecast for the year made prior to the earnings announcement date is negative, 0 otherwise. We estimate the expected part of the forecast based on prior-year earnings and cumulative returns during the year. Appendix A defines the remaining variables. All observations outside the interval –£0.02 ≤ *SURP* < £0.02 are deleted when *JMBE* is used, resulting in 868 firms and 2,768 observations. *SURP* is the earnings surprise measured as the difference between I/B/E/S actual EPS and the latest analyst forecast made prior to the earnings announcement date.

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# The effects of voluntary disclosure and dividend propensity on prices leading earnings

**Khaled Hussainey and Martin Walker***

**Abstract**—We investigate the joint effects of dividend propensity (i.e. whether a firm pays cash dividends) and voluntary disclosure on the relationship between current stock returns and future earnings. We examine whether dividend propensity and voluntary disclosure act as substitutes or complements in the financial communication process. We also examine whether the effects of dividend propensity and voluntary disclosure vary between high- and low-growth firms.

Consistent with prior studies, we find that share price anticipation of earnings improves with increasing levels of annual report narrative disclosure, and that firms that pay dividends exhibit higher levels of share price anticipation of earnings than non-dividend-paying firms.

The paper adds to the literature on share price anticipation of earnings in two crucial respects. First we show that the associations of voluntary disclosure and dividend propensity with share price anticipation of earnings are statistically significant for high-growth firms and insignificant for low-growth firms. Second we show that the significant effects we find for dividend propensity and voluntary disclosure in high-growth firms are not perfectly additive.

**Keywords:** annual report narratives; dividend propensity; firm growth characteristics; returns-earnings relationship; content analysis

## 1. Introduction

Considerable attention has been given to examining the association between corporate disclosure and share price anticipation of earnings (e.g. Schleicher and Walker, 1999; Lundholm and Myers, 2002; Gelb and Zarowin, 2002; Hussainey et al., 2003 and Schleicher et al., 2007). These papers find that the stock market's ability to anticipate future earnings changes is significantly improved when firms voluntarily provide higher levels of disclosure. However, these studies do not take into account the possibility that dividend policy may provide an alternative device for conveying value-relevant information to the stock market that might act as a substitute or complement for narrative disclosure in the financial communication process.

Hanlon et al. (2007) examine the impact of dividend propensity (i.e. whether a firm pays cash dividends) on the stock market's ability to anticipate future earnings. They modify and use the returns-earnings regression model introduced by Collins et al. (1994) to compare the association between current-year stock returns and future earnings for firms that pay dividends in the current year as compared with non-dividend-paying firms. They find that US dividend-paying firms exhibit significantly higher levels of share price anticipation of earnings than non-dividend-paying firms. In addition, Hanlon et al. (2007) control for disclosure quality, as measured by AIMR-FAF scores, and they find that the significance of the dividend policy for anticipating future earnings is reduced. This suggests that dividends and disclosure might be substitute forms of financial communication. However, it is possible that the relative information content of dividends and voluntary disclosure could be different in the US than in the UK as the proportion of UK dividend-paying firms is greater than US dividend-paying firms in the period of 1996–2002 (73% in the UK compared with 23% in the US; see Denis and Osobov, 2008 for more details).[1]

The present paper examines the joint effects of

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[1] One possible reason for this is the difference between the treatments of dividend income in the two countries (Morgan and Thomas, 1998). In particular, the UK imputation system up to 5 April 1997 significantly favoured dividend payments compared to the US for both individuals and pension funds and thereafter still favoured dividend payments to individuals. The abolition of advance corporation tax in April 1999 did not offer this (Bank et al., 2006). Oswald and Young (2008: 51–52) provide detailed information on dividends and relevant UK tax legislation.

dividend propensity and voluntary disclosure on share price anticipation of earnings.

In undertaking this task, we argue that it is vital to take into account the growth characteristics of firms. There are strong theoretical and empirical grounds for expecting this to be the case. Several researchers (see, for example, Brown et al., 1999; Francis and Schipper, 1999 and Lev, 1989) have identified a number of problems with the financial reporting process, instances of accounting 'failure'. Particular attention has been paid to the inability of the financial reporting system to capture the value-relevance of intangible investments on a timely basis (see, for example, Amir and Lev, 1996; Lev, 2001 and Lev and Sougiannis, 1996). High growth and intangible asset intensity are factors that tend to reduce the predictive value of current earnings for future earnings. Investors of high-growth firms are aware that current earnings provide a poor guide to the future financial performance of the firm. Thus, when valuing these firms, investors tend to seek other, more timely, predictors of future earnings beyond current earnings.

In this paper we use the future earnings response model of Collins et al. (1994) to measure the degree of share price anticipation of earnings. We test whether the level of share price anticipation of earnings varies with dividend propensity and with the level of voluntary disclosure. We also test whether the associations between dividend propensity or voluntary disclosure and share price anticipation of earnings differ between high- and low-growth firms.

Our results show that both narrative disclosure level and dividend propensity are associated with significantly improved share price anticipation of earnings for high-growth firms. Moreover, for high-growth firms, narrative disclosure and dividend propensity appear, to some extent, to be substitute forms of financial communication. In contrast we find that neither narrative disclosure level nor dividend propensity exhibit significant association with share price anticipation of earnings for low-growth firms.

Thus the paper makes an important and novel contribution to the literature on corporate financial communication. So far as we are aware, it is the only paper to examine the joint role of narrative disclosure and dividend propensity on prices leading earnings. Moreover, it is the first paper to show that the predictive value of dividend propensity and narrative disclosure is sensitive to the growth characteristics of the firm.

The paper proceeds as follows. Section 2 reviews the prior literature and develops our hypotheses. In Section 3, we describe the generation of our disclosures scores. Our measure of share price anticipation of earnings is discussed in Section 4. Section 5 describes the data and we present our main regression results in Section 6. Section 7 presents our specification check and Section 8 concludes and suggests areas for future research.

## 2. Prior research and hypotheses development

### *2.1. Disclosure and prices leading earnings*

There is a growing body of literature which examines how corporate disclosures affect the stock market's ability to anticipate future earnings changes. Hussainey et al. (2003) examine the information content of annual report narrative sections for anticipating future earnings for UK firms. Their work builds on the earlier work of Schleicher and Walker (1999) but, by adopting the augmented returns-earnings regression model of Collins et al. (1994), uses a research design closer to those used in Gelb and Zarowin (2002) and Lundholm and Myers (2002).

Hussainey et al. (2003) and Schleicher and Walker (1999) find that improved levels of annual report disclosures tend to lead to higher levels of share price anticipation of earnings. In particular, Hussainey et al. (2003) find that forward-looking earnings statements in the annual report narratives increase the market's ability to anticipate future earnings change. However, they do not find significant results when using a disclosure metric based on all types of forward-looking statements. Gelb and Zarowin (2002) and Lundholm and Myers (2002) find that the quality of corporate disclosure, as measured by AIMR-FAF analysts' rankings of disclosure, is positively associated with the market's ability to anticipate future earnings changes. Schleicher et al. (2007) find that the association between levels of annual report disclosures and share price anticipation of earnings is not the same for profitable and unprofitable firms. They find that the ability of stock returns to anticipate next period's earnings change is significantly greater for unprofitable firms that provide higher levels of earnings predictions in their annual report narratives. They did not find such results for profitable firms.

None of the above papers makes any distinction between high-growth firms and low-growth firms. However all of these studies make use of the relationship between current and future earnings and stock returns to assess the information content of corporate disclosures. Moreover, it is well known that differences in growth rates cause the association between stock returns and contemporaneous earnings changes to vary between firms (Collins and Kothari, 1989 and Rayburn, 1986). High-growth firms tend to have higher levels of intangible assets (Thornhill and Gellatly, 2005). Such

intangible assets tend to reduce the value-relevance of current earnings for overall firm value. If the benefits from these assets are uncertain, investors will find it more difficult to appraise firm value (Kothari et al., 2002). Additionally, Lee et al. (2005) argue that increasing uncertainty about future benefits will lead to more information asymmetry between investors and managers and may introduce noise in the estimation of firm value in high-tech in comparison with low-tech firms.[2] Firms can reduce such information asymmetry by providing additional voluntary information (Ertimur, 2004). In this paper, we ask whether increasing the level of annual report forward-looking disclosures improves the stock market's ability to anticipate future earnings changes, especially for high-growth firms.

While no papers have examined the effect of firm-specific growth characteristics on the association between disclosure and the stock market's ability to anticipate future earnings changes, possibly the most closely related paper to ours is Kwon (2002). Kwon (2002) compares analysts' forecast accuracy and dispersion between high-tech and low-tech firms. He finds that high-tech firms have lower error and dispersion of analyst earnings forecasts than low-tech firms. In addition, he finds that high-tech firms have higher analyst forecast accuracy than low-tech firms. He attributes this finding to differences in the level of voluntary disclosure provided by these firms. In particular, he states that high-tech firms increase their levels of voluntary disclosure to attract more analysts. Increasing disclosure leads to higher levels of analyst forecast accuracy and lower levels of forecast dispersion. The finding of Kwon's paper is important in relation to ours as it indicates that firm characteristics such as growth prospects can condition the forecasting practices of stock market participants.

Considerable attention has been paid to the inability of financial performance measures to capture the value relevance of intangible investments on a timely basis (see Abdolmohammadi et al., 2006; Lee et al, 2005; Lev, 2001; Lev and Zarowin, 1999). As a consequence, the predictive value of current earnings for future earnings is lower for intangibles-rich high-growth firms than for low-growth firms.

Given the limited usefulness of their current earnings for predicting future earnings, one possible response for firms with significant unrecognised intangible assets or high-growth prospects is to make voluntary disclosure in order to convey value-relevant information in a more timely way (Gelb, 2002; Hyytinen and Pajarinen, 2005 and Khurana et al., 2006). Moreover, since current earnings provide a better basis for predicting future earnings for low-growth firms, we expect the influence of disclosure on the relationship between current returns and future earnings to be stronger for high-growth firms than low-growth firms.

The following hypotheses focus on how voluntary disclosure affects the stock market's ability to anticipate future earnings:

**H1** The degree of share price anticipation of earnings is positively related to the level of voluntary disclosure.

**H2** The degree of share price anticipation of earnings is positively related to the level of voluntary disclosure for high-growth firms.

**H3** The degree of share price anticipation of earnings is positively related to the level of voluntary disclosure for low-growth firms.

We also compare the effect of voluntary disclosure between high- and low-growth firms. Thus, the fourth hypothesis states:

**H4** The strength of the degree of association between share price anticipation of earnings and voluntary disclosure is greater for high-growth firms than for low-growth firms.

### *2.2. Dividends and prices leading earnings*

Apart from the audited financial statements, voluntary disclosure is one of two main ways that firms can communicate information about future earnings to the stock market. The other way is through some types of financial policy choices.

There is a vast literature that explores the possibility that financial policy choices may serve to convey information to the market. One such line of argument focuses on financial signalling as a solution to the adverse selection problem. In this paper we focus on one particular form of financial signalling, cash dividends. There is a substantial amount of literature on dividend signalling in its various forms. This literature has produced a complex set of models that can be used to rationalise alternative types of dividend signalling behaviour, that can represent different forms of dividend policy, and that generate different models of the relationship between company cash flows, dividends, and share prices (see Allen and Michaely, 2003; Benartzi et al., 1997; Eades, 1982; and Grullon et al., 2003 for surveys).

Another line of argument focuses on financial policy choices as a part of the solution to investor/manager agency conflicts. For example, Jensen (1986) points to the potential agency costs of firms having very high free cash flows, and the need to set limits to the discretionary investment

[2] High-tech firms invest more in unrecognised intangible assets (Kwon, 2002). As a consequence, our high-growth sample contains a much larger proportion of high-tech firms than the low-growth sample.

choices of company managers. Under this line of reasoning, the payment of dividends is potentially informative about the quality of external investor protection in the firm. Note that under this line of reasoning, the communication of information is incidental to the governance role of dividends.

Some recent studies explore the nature of the information revealed by dividends (i.e. Grullon et al., 2002 and Nam et al., 2008). These studies suggest that dividends changes are associated with changes in firm risk. In particular, Nam et al. (2008) show that firms that initiate dividend payment experience a reduction in risk, and Grullon et al. (2002) show that firms that increase their payout ratios experience reduced risk.

A number of papers have studied the empirical links between earnings quality (in particular earnings persistence) and dividend payouts. The recent papers of Garcia-Borbolla et al. (2004) and Skinner (2004) identify subtle and complex interactions between dividend propensity and the quality of earnings. Moreover, these studies reveal significant differences between US and European firms. Skinner's study suggests that dividends are more likely to be paid by the larger, more stable US firms. Such firms tend to have more predictable earnings streams. Thus his work suggests that the predictability of earnings from current earnings and the payment of dividends are mutually related through the underlying stability of the firm. On the other hand, the study by Garcia-Borbolla et al. (2004) of European firms concludes that dividends are more useful for predicting future earnings where the quality of earnings is low.

This paper makes no attempt to identify the specific types of information that dividend payments convey to investors. As in Hanlon et al. (2007) we simply start from the observation that not all firms pay dividends. Firms that do not pay dividends, by definition, cannot be using dividends to communicate value-relevant information. On the other hand, the payment of dividends may change the ability of the market to anticipate future earnings changes.

Relative to Hanlon et al. we make three contributions. First we present results for the UK economy for a period in which the propensity to pay dividends was much higher than in the US. The average dividend propensity for the firms in our sample is 84%, considerably higher than the corresponding propensity for US firms. For example, Skinner (2008) reports an average dividend propensity of 28% for US listed firms for the years 1995–2004. Second, we jointly model the effects of voluntary disclosure and dividend propensity on share price anticipation of earnings. Third, we model the extent to which the effects of dividend propensity and voluntary disclosure on share price anticipation of earnings are linked to the growth characteristics of firms.

Hanlon et al. (2007) make no distinction between high- and low-growth firms. However it can be argued that the importance of other information for predicting future earnings could differ between high- and low-growth situations. High-growth firms typically exhibit higher levels of information asymmetry than low-growth firms, and high-growth firms are more likely to need to raise external capital in order to finance their dividend payments whilst maintaining high levels of investments. Thus there are good reasons to expect the effects of voluntary disclosure and dividend propensity to vary between high- and low-growth firms.

Thus we state the following hypotheses:

**H5** The degree of share price anticipation of earnings is greater for dividend-paying firms than for non-dividend-paying firms.

**H6** The degree of share price anticipation of earnings is greater for dividend-paying high-growth firms than non-dividend-paying high-growth firms.

**H7** The degree of share price anticipation of earnings is greater for dividend-paying low-growth firms than non-dividend-paying low-growth firms.

**H8** The strength of the degree of association between share price anticipation of earnings and dividend propensity is greater for high-growth firms than low-growth firms.

### *2.3. The joint effect of disclosure and dividends on prices leading earnings*

Having introduced the main hypotheses relating to share price anticipation of earnings and the two forms of financial communication we now consider how to test whether the two forms of communication act as complements or substitutes. There are four logical possibilities:

First, if voluntary disclosure and dividend payments are different ways of conveying the same information, then firms that have high levels of disclosure but pay no dividends, should exhibit roughly the same degree of share price anticipation of earnings as firms with high levels of disclosure that pay dividends. Similarly firms that pay dividends should have roughly the same level of share price anticipation of earnings irrespective of their level of disclosure.

Second, if dividend payments and voluntary disclosure convey unrelated types of information, then the level of share price anticipation of earnings for firms that have high levels of disclosure and pay dividends should be stronger than the level of share price anticipation of earnings for firms that have high levels of disclosure but do not

pay dividends. Similarly, the level of share price anticipation of earnings should be stronger when both types of communication are present than when only the payment of dividends is present.

Third, if the combination of dividend payments and voluntary disclosure produces related information that is 'reinforcing' (i.e. if there is a multiplicative effect) then share price anticipation of earnings will be the greatest for firms that have high disclosure and also pay dividends.

Finally, if dividend payments and voluntary disclosure convey related information, but some of the information is common to both, i.e. 'partially additive', then the level of share price anticipation of earnings for firms that have high levels of disclosure and pay dividends should be higher than the level of share price anticipation of earnings when firms have high levels of disclosure but do not pay dividends or the level of share price anticipation of earnings should be higher when both types of communications are present than when only the payment of dividends is present. We test to see which of these four possibilities is present in the data by allowing for an interactive effect in our model. Thus we state the following hypotheses:

**H9** The effects of voluntary disclosure and dividend propensity on the degree of share price anticipation of earnings are additive.

**H10** The effects of voluntary disclosure and dividend propensity on the degree of share price anticipation of earnings for high-growth firms are additive.

**H11** The effects of voluntary disclosure and dividend propensity on the degree of share price anticipation of earnings for low-growth firms are additive.

**H12** The strength of the joint effect of disclosure and dividend propensity on share price anticipation of earnings is the same for high- and low-growth firms.

To test the above hypotheses, we follow the approach of Hussainey et al. (2003) to automate the generation of forward-looking earnings disclosure scores. Section 3 provides further details. We use the modified regression model of Collins et al. (1994) to measure the influence of voluntary disclosure and dividend propensity on share price anticipation of earnings. This model is discussed in Section 4.

## 3. Disclosure scores

We adopt the scoring methodology developed in Hussainey et al. (2003). They automate the generation of disclosure scores for large samples of UK firms through the use of QSR N6, a text analysis software package.[3] We focus on annual report narratives because they are more likely to contain voluntary forward-looking earnings predictions than other sections of the annual report. We choose narrative sections with at least one of the following headings: Financial Highlights, Summary Results, Chairman's Statement, Chief Executive Officer's Review, Operating and Financial Review, Financial Review, Financial Director's Report, Finance Review, Business Review, and Operating Review. All other sections of the annual report are excluded from our analysis (for more details, see Hussainey et al., 2003).

To measure the informativeness of a firm's narrative disclosures, we identify the forward-looking earnings sentences that are most likely to be useful for predicting a firm's future earnings changes. Our measure of disclosure quality is the number of forward-looking sentences in the annual report narratives that contain earnings indicators. We focus on earnings indicators because Hussainey et al. (2003) and Schleicher et al. (2007) find that these indicators improve the stock market's ability to anticipate future earnings change one year ahead.

We calculate our disclosure scores in three steps. The first step requires the identification of all sentences that are associated with forward-looking statements in annual report narratives. In this step, we text-search the narrative sections of annual reports using the list of forward-looking key words adopted in Hussainey et al. (2003: 277). This list includes the following 35 keywords: accelerate, anticipate, await, coming (financial) year(s), coming months, confidence (or confident), convince, (current) financial year, envisage, estimate, eventual, expect, forecast, forthcoming, hope, intend (or intention), likely (or unlikely), look forward (or look ahead), next, novel, optimistic, outlook, planned (or planning), predict, prospect, remain, renew, scope for (or scope to), shall, shortly, should, soon, will, well placed (or well positioned), year(s) ahead. Similar to Hussainey et al. (2003) we also include future year numbers in the list of forward-looking keywords.

The next step in the generation of earnings disclosure scores is the identification of information items that are relevant to the capital market in assessing the firm's future earnings. Since the capital market's information set is unobservable, Hussainey et al. (2003) examine the contents of sell-side analysts' reports as a proxy for the market's view about the firm's disclosure quality. For each forward-looking statement in analysts' reports, they identify the key noun of that statement. For the purpose of the current paper, we use the

[3] In the current paper, we use QSR N6 to further facilitate the automation of text searches. Further information about QSR N6 is available online at http://www.qsrinternational.com.

same list adopted in Hussainey et al. (2003: 280) that is related to earnings indicators. The list contains the following 12 keywords: benefit, break even, budget, contribution, earnings, EPS, loss, margin, profit, profitability, return and trading.

Finally, we use QSR N6 to count the number of sentences that include both at least one forward-looking keyword and at least one earnings indicator. This is done by finding the intersection of the keyword search and the topic search.

## 4. A measure of prices leading earnings

Our measure of prices leading earnings is based on Collins et al. (1994). They use the future earnings response coefficient (FERC) as a proxy for the stock market's ability to anticipate future earnings. FERC is estimated by regressing current-year stock returns on current and future annual earnings and returns plus control variables. The regression model of Collins et al. (1994) has been applied in a large number of recent papers, e.g. Banghøj and Plenborg, 2008; Dhiensiri et al., 2005; Ettredge et al., 2005; Gelb and Zarowin, 2002; Hanlon et al., 2007; Hussainey et al., 2003; Lee et al., 2007; Lundholm and Myers, 2002; Orpurt and Zang, 2006; Oswald and Zarowin, 2007; Schleicher et al., 2007 and Tucker and Zarowin, 2006. In effect the regression model of Collins et al. (1994) has become the standard technique for measuring prices leading earnings.

Collins et al. (1994) start by highlighting the poor empirical performance of the basic contemporaneous returns earnings regression.

$$R_t = b_0 + b_1 X_t + u_t \quad (1)$$

Where $R_t$ is the stock return for year $t$. $X_t$ is defined as earnings change deflated by price at $t$–1. Under ideal conditions[4] equation (1) will yield a perfect fit, and the earnings response coefficient (ERC) will be equal to $\frac{1+r}{r}$, where $r$ is the required rate of return on equity (Walker, 2004). Numerous attempts to estimate equation (1) on annual data have revealed a very poor statistical fit (an $R^2$ of 10% or less) and an ERC between 1 and 3, i.e. much lower than the value implied by a typical cost of equity capital.

Collins et al. (1994) argue that current returns reflect information about both current and future earnings. Therefore any attempt to explain current stock return in terms of earnings changes, should control for information about future earnings received in the current period. In the light of this argument, they include three future earnings growth variables (N=3 and k=1, 2, 3) and make a number of adjustments to equation (1) to arrive at the following regression model:[5]

$$R_t = b_0 + b_1 X_t + \sum_{k=1}^{N} b_{k+1} X_{t+k} + \sum_{k=1}^{N} b_{k+N+1} R_{t+k} + b_{2N+2} AG_t + b_{2N+3} EP_{t-1} + e_t \quad (2)$$

where:

$R_t$ is the stock return for year $t$.

$R_{t+1}$, $R_{t+2}$ and $R_{t+3}$ are the stock returns of years $t$+1, $t$+2 and $t$+3 respectively.

$X_t$, $X_{t+1}$, $X_{t+2}$ and $X_{t+3}$ are defined as earnings change deflated by lagged earnings at $t$–1.

$EP_{t-1}$ is earnings of period $t$–1 over price starting four months after the financial year-end of period $t$–1.

$AG_t$ is the growth rate of total book value of assets for period $t$.

Equation (2) shows that a number of forward-dated variables are introduced in order to measure the level of prices leading earnings. Specifically, the model includes future earnings changes as proxies for the information received by the market in period $t$ about earnings growth in years $t$+1 and beyond. The inclusion of forward returns in the model, $R_{t+1}$, $R_{t+2}$ and $R_{t+3}$, controls for news about earnings growth in period $t$+1, $t$+2 and $t$+3 received in period $t$+1, $t$+2 and $t$+3 respectively. Because the forward returns variables control for news received in the future about future earnings, the forward earnings variables $X_{t+1}$, $X_{t+2}$ and $X_{t+3}$ proxy for news about future earnings received in the current period, $t$. This model includes the contemporaneous asset growth rate, $AG_t$ to control for the possibility that firms may invest in advance of future earnings. It also includes the earnings-price ratio, $EP_{t-1}$, to control for the possibility that the earnings of period $t$–1 are not a good proxy for the market's expectations (at time $t$–1) of the earnings for period $t$ and beyond.

For the ease of exposition we following the idea of Lundholm and Myers (2002) and Oswald and Zarowin (2007) in aggregating future earnings over three years into one future variable, $X_{t3}$ and future returns over three years into one future variable, $R_{t3}$.[6] Such aggregations produce the following regression model:

$$R_t = b_0 + b_1 X_t + b_2 X_{t3} + b_3 R_{t3} + b_4 AG_t + b_5 E_{t-1} + e_t \quad (3)$$

[4] Walker (2004) argues that ideal conditions require semi-strong market efficiency. In such a market, earnings follow a random walk and earnings at time $t$ capture all value-relevant information available at time $t$.

[5] Note that Collins et al. (1994) find that the association between current stock returns and future earnings is not significant beyond three years ahead.

[6] Lundholm and Myers (2002) show that their results are unchanged whether the three future years are aggregated or separated.

To test our main hypotheses, we use the modified version of Collins et al. (1994), equation (3). However, similar to Hussainey et al. (2003) and Schleicher et al. (2007), in defining the earnings growth variable, we deflate earnings change by the share price at the start of the current year and not by lagged earnings. This is due to the fact that it is difficult to define earnings growth when lagged earnings are negative or zero. As a result, a price deflator is used instead of the earnings deflator.

As argued earlier, current earnings numbers are likely to be less useful for predicting earnings for high-growth firms. So we predict that these firms will use other indicators such as voluntary disclosure or dividend decisions to convey value-relevant information in addition to current earnings. Such information should enable the market to better anticipate a firm's future earnings. This should lead to high voluntary disclosure firms or dividend-paying firms having a stronger relationship between current returns and future earnings changes than low-disclosure firms or non-dividend-paying firms. Therefore, we predict higher FERCs for high-disclosure firms or dividend-paying firms.

To test this prediction we interact all independent variables in equation (3) with a dummy variable, $D$, defined to be one for high-disclosure firms and zero otherwise. We do not use the actual disclosure scores. Instead we define $D$ to be 1 for firms in the top two quartiles of the distribution of disclosure scores and 0 otherwise.[7] We also interact all independent variables with a dummy variable, $Div$, defined to be one for firms that pay a dividend in the current year and zero otherwise. Finally we extend the model to test for the interaction between $D$ and $Div$. Interacting all explanatory variables in (3) with $D$, $Div$, and $D*Div$ yields our main regression model:

$$R_t = b_0 + b_1X_t + b_2X_{t3} + b_3R_{t3} + b_4AG_t + b_5EP_{t-1} \quad (4)$$
$$b_6D + b_7D^*X_t + b_8D^*X_{t3} + b_9D^*R_{t3} + b_{10}D^*AG_t + b_{11}D^*EP_{t-1}$$
$$b_{12}Div + b_{13}Div^*X_t + b_{14}Div^*X_{t3} + b_{15}Div^*R_{t3} + b_{16}Div^*AG_t + b_{17}Div^*EP_{t-1}$$
$$b_{18}D^*Div + b_{19}D^*Div^*X_t + b_{20}D^*Div^*X_{t3} + b_{21}D^*Div^*R_{t3} + b_{22}D^*Div^*AG_t + b_{23}D^*Div^*EP_{t-1} + e_t$$

The coefficient on $X_t$ is hypothesised to be positive. The coefficient on $X_{t3}$ measures three years ahead share price anticipation of earnings for non-dividend-paying firms with low disclosure scores. This is the base case in the model. All independent variables are interacted with both of the two dummy variables $D$ and $Div$. The coefficient on $D*X_{t3}$ measures the extent to which share price anticipation of earnings is greater for high-disclosure non-dividend-paying firms compared to the base case (i.e. it measures the pure disclosure effect). The coefficient on $Div*X_{t3}$ measures the extent to which share price anticipation of earnings is greater for low-disclosure dividend-paying firms compared to the base case.

We expect the regression coefficients associated with $D*X_{t3}$ and $Div*X_{t3}$ to be significantly positive for high-growth firms. In addition we expect these coefficients to be smaller for low-growth firms than for high-growth firms.

The variable $D*Div*X_{t3}$ measures the incremental effect of both high disclosure and dividend propensity. There are four logical possibilities:[8]

First, both disclosure and dividends provide the same information. In this case, the coefficient on $D*X_{t3}$ will be equal to the coefficient on $Div*X_{t3}$. In addition, the coefficient on $D*Div*X_{t3}$ should be negative and equal in absolute value to the coefficients on $D*X_{t3}$ or $Div*X_{t3}$. As a result, the total impact of both disclosure and dividends should be calculated as follows:

$$D^*X_{t3} + Div^*X_{t3} + D^*Div^*X_{t3} = D^*X_{t3}$$
$$(\text{or} = Div^*X_{t3}, \text{ since } D^*X_{t3} = Div^*X_{t3}).$$

Second, both types of communication provide ('additive') unrelated information. In this case, we predict the coefficient on $D*Div*X_{t3}$ should not be significantly different from zero. As a result, the total impact of both disclosure and dividends should be calculated as follows:

$$D^*X_{t3} + Div^*X_{t3} - D^*Div^*X_{t3} = D^*X_{t3} + Div^*X_{t3}$$

Third, both types of communication provide related information that is 'reinforcing' or 'multiplicative'. In this case, the coefficient on $D*Div*X_{t3}$ should be significantly greater than zero. In other words, the sum of the coefficients on $D*X_{t3}$, $Div*X_{t3}$ and $D*Div*X_{t3}$ should be significantly greater than the sum of the coefficients on $D*X_{t3}$ and $Div*X_{t3}$. In this case, the inference is that both dividend propensity and voluntary disclosure are strictly complementary.

Finally, both types of communication provide related information, but some of the information is common to both, i.e. 'partially additive'.

[7] In Hussainey et al. (2003) and Gelb and Zarowin (2002), the authors drop observations with disclosure scores in the second and third quartiles. However, we use the full sample without dropping observations in the middle quartiles to maintain a usable sample size for the regression analyses. As we will discuss later, Table 1, Panel C shows that the number of usable non-dividend-paying firms is 551. This number comprises 297 high-growth firm-years and 254 low-growth firm-years. Deleting firms in the middle quartiles will significantly reduce the usable number of observations in each growth category (e.g. the number will be 142 for high-growth firms and 111 for low-growth firms).

[8] The authors thank an anonymous referee for the suggestion.

Therefore, we predict that the coefficient on $D^*Div^*X_{t3}$ should be significantly lesser than zero. In other words, the sum of the coefficients on $D^*X_{t3}$, $Div^*X_{t3}$ and $D^*Div^*X_{t3}$ should be significantly less than the sum of the coefficients on $D^*X_{t3}$ and $Div^*X_{t3}$. In this case, the inference is that both dividend propensity and voluntary disclosure are partial substitutes.

We treat the issue of which of these four logical possibilities is true as a purely empirical question and offer no prior theoretical predictions as to which of these outcomes is the most likely.

## 5. Data

To perform our analyses, annual reports in an electronic format are required in order to use QSR N6 software as a text analysis tool. Therefore, our initial sample is limited to all UK non-financial firms on the *Dialog* database that have at least one annual report. *Dialog* covers large cross-sections of electronic non-financial UK annual reports for the years 1996–2002. So, we limit our study to that sample period.[9] We use book-to-market value as a measure of growth. We measure growth on a yearly basis. This allows us to examine the effect of disclosure and dividends on share price anticipation of earnings when firms are classified as high-growth or low-growth in a particular time period. We identify high-growth firms as those having below median levels of book-to-market value and low-growth firms as those having above median levels of book-to-market value.

The total number of annual reports on *Dialog* for non-financial firms for the sample period is 8,098. Of those, 7,977 firm-years have matching records in *Datastream*. We delete 1,312 firm-years observations because of changing year-ends. This leaves 6,665 firm-years. We also delete 2,958 firm-years missing observations. This leaves 3,707 firm-year observations. Finally, we delete outliers defined as observations falling into the top or bottom 1% of the distribution of any of the regression variables. Following Schleicher et al. (2007), we treat the observations of high- and low-growth firms as separate distributions. Otherwise, an unreasonably large number of high-growth firms' observations will fall into the top and bottom 1%. Deletion of inappropriate observations and observations with missing data reduces the sample to 3,503 firm-years observations. Of those, 1,770 firm-years are high-growth firms and 1,733 firm-years are low-growth firms.

Earnings per share data is calculated by dividing operating income before all exceptional items (*Worldscope* item 01250) by the outstanding number of shares. $X_t$, $X_{t+1}$, $X_{t+2}$ and $X_{t+3}$ are then defined as the earnings change for the periods $t$, $t$+1, $t$+2 and $t$+3 deflated by the share price. Both current and future earnings changes are deflated by the share price at the start of the return window for period $t$. $X_{t3}$ is calculated as the aggregated future earnings over the following three years relative to the financial year-end. We collect returns data from *Datastream*. $R_t$, $R_{t+1}$, $R_{t+2}$ and $R_{t+3}$ are measured as buy-and-hold returns starting from eight months before the financial year-end to four months after the financial year-end in year $t$. In the return measures, similar to Hussainey et al. (2003), we incorporate a four-month lag to ensure that annual reports have been released. $R_{t3}$ is calculated as the aggregated future returns over the following three years relative to the financial year-end. Earnings yield, $EP_{t-1}$, is defined as period $t$–1's earnings over price four months after the financial year-end of period $t$–1. $AG_t$ is the growth rate of total book value of assets for period $t$ (*Datastream* item 392).

We collect dividends per share from *Worldscope* (item 05101).[10] The dividends dummy variable, $Div$, is set equal to one if firms pay dividends in year $t$ and zero otherwise. In addition to the above variables, we use a disclosure dummy variable to examine the effect of disclosure on the returns-earnings association. The disclosure dummy, $D$, is set equal to one for firms in the top two quartiles of the distributions of disclosure scores and zero otherwise. As mentioned earlier, we identify high- (low-) growth firms as those having below (above) median levels of BTMV. *BTMV* ratio is calculated as the inverse of the market to book value of equity ratio (*Datastream* item: *MTBV*).

Table 1 presents descriptive statistics for our regression variables. Panel A reports the descriptive analysis for the full sample. Panel B (C) reports the descriptive analysis for dividend (non-dividend) paying firms. Panel D (E) reports the descriptive analysis for high- (low-) growth firms.

Table 1 shows that dividends were paid in 84% (=2952/3503) of the firm years in our sample.[11]

[9] Dividends are usually tested over longer sample periods. However, we restrict our analyses to the years in which large numbers of annual reports are available on *Dialog*. This enables us to test the joint effect of disclosure and dividends on prices leading earnings.

[10] *Worldscope* defines dividends per share (item 05101) as 'the total dividends per share declared during the calendar year for U.S. corporations and fiscal year for Non-US corporations. It includes extra dividends declared during the year. Dividends per share are based on the 'gross' dividend of a security, before normal withholding tax is deducted at a country's basic rate, but excluding the special tax credit available in some countries'.

[11] The proportion of dividend-paying firms for the overall population of UK firms is 69%. This number is calculated by dividing the total number of dividend-paying firms by the total number of firms in the period of 1996–2002. This number is consistent with a recent study by Denis and Osobov (2008) which find that 73% of UK firms pay dividends in the sample period 1996–2002. However, in our sample, the proportion of dividend-paying firms is much higher than the overall population of dividend-paying firms. This is due to the deletion of inappropriate observations and observations with missing data in the sample period of 1996–2002.

**Table 1**
**Descriptive statistics**

| *Variable* | *Mean* | *Median* | *OBS* |
|---|---|---|---|
| **Panel A: Full sample** | | | |
| $R_t$ | 0.089 | 0.038 | 3503 |
| $X_t$ | 0.007 | 0.008 | 3503 |
| $X_{t3}$ | 0.006 | 0.005 | 3503 |
| $R_{t3}$ | 0.295 | 0.222 | 3503 |
| $AG_t$ | 0.173 | 0.078 | 3503 |
| $EP_{t-1}$ | 0.091 | 0.093 | 3503 |
| Disclosure = Low | 1.7 | 2 | 1674 |
| Disclosure = High | 7.2 | 6 | 1829 |
| **Panel B: Dividend-paying firms** | | | |
| $R_t$ | 0.116 | 0.065 | 2952 |
| $X_t$ | 0.006 | 0.008 | 2952 |
| $X_{t3}$ | –0.002 | 0.003 | 2952 |
| $R_{t3}$ | 0.301 | 0.237 | 2952 |
| $AG_t$ | 0.171 | 0.085 | 2952 |
| $EP_{t-1}$ | 0.119 | 0.102 | 2952 |
| Disclosure = Low | 1.8 | 2 | 1297 |
| Disclosure = High | 7.4 | 6 | 1655 |
| **Panel C: Non-dividend-paying firms** | | | |
| $R_t$ | –0.055 | –0.187 | 551 |
| $X_t$ | 0.017 | 0.001 | 551 |
| $X_{t3}$ | 0.050 | 0.024 | 551 |
| $R_{t3}$ | 0.267 | 0.082 | 551 |
| $AG_t$ | 0.184 | 0.015 | 551 |
| $EP_{t-1}$ | –0.058 | –0.020 | 551 |
| Disclosure = Low | 1.4 | 1 | 377 |
| Disclosure = High | 6.1 | 5 | 174 |
| **Panel D: High-growth firms** | | | |
| $R_t$ | 0.089 | 0.033 | 1770 |
| $X_t$ | 0.007 | 0.008 | 1770 |
| $X_{t3}$ | 0.002 | 0.003 | 1770 |
| $R_{t3}$ | 0.234 | 0.166 | 1770 |
| $AG_t$ | 0.220 | 0.114 | 1770 |
| $EP_{t-1}$ | 0.073 | 0.079 | 1770 |
| Disclosure = Low | 1.7 | 2 | 828 |
| Disclosure = High | 7.3 | 6 | 942 |
| **Panel E: Low-growth firms** | | | |
| $R_t$ | 0.089 | 0.043 | 1733 |
| $X_t$ | 0.008 | 0.007 | 1733 |
| $X_{t3}$ | 0.010 | 0.007 | 1733 |
| $R_{t3}$ | 0.358 | 0.284 | 1733 |
| $AG_t$ | 0.124 | 0.050 | 1733 |
| $EP_{t-1}$ | 0.110 | 0.116 | 1733 |
| Disclosure = Low | 1.7 | 2 | 846 |
| Disclosure = High | 7.1 | 6 | 887 |

**Table 1**
**Descriptive statistics** *(continued)*

Table 1 presents descriptive statistics. Panel A reports the descriptive analysis for the full sample. Panel B (C) reports the descriptive analysis for dividend (non-dividend) paying firms. Panel D (E) reports the descriptive analysis for high- (low-) growth firms. Returns, $R_t$, is calculated as buy-and-hold returns from eight months before the financial year-end to four months after the financial year-end. $R_{t3}$ is the aggregated three years future returns. The earnings variable, $X_t$, is defined as earnings change per share deflated by the share price four months after the end of the financial year $t$–1. $X_{t3}$ is the aggregated three years future earnings change. Earnings measure is the *Worldscope* item 01250 which is operating income before all exceptional items. $EP_{t-1}$ is defined as period $t$–1's earnings over price four months after the financial year-end of period $t$–1. $AG_t$ is the growth rate of total book value of assets for period $t$ (*Datastream* item 392). Firm-years with a disclosure score on top (bottom) 50% of the distribution of disclosure scores are defined as high- (low-) disclosure firm-year. Dividend (non-dividend) paying firms are those paying (not paying) dividends at the current year. Dividends measure is dividends per share, *Worldscope* item 05101. High-growth firms are defined as those having below median levels of BTMV, while low-growth firms are defined as those having above median levels of BTMV. *BTMV* ratio is calculated as the inverse of the market-to-book value of equity ratio (*Datastream* item *MTBV*).

Panels B and C of Table 1 show that the median return in period $t$ is positive for dividend-paying firms, whilst it is negative for non-dividend-paying firms. The median current and future earnings change is positive for both dividend-paying and non-dividend-paying firms.

Panels B and C of Table 1 also show that dividend-paying firms, on average, have slightly higher mean and median levels of voluntary disclosure than non-dividend-paying firms. 56% of dividend-paying firms are in the category of high-disclosure firms compared to 32% of non-dividend-paying firms.[12] The table shows a material difference in the disclosure scores between firms in high- and low-disclosure categories for both dividend and non-dividend-paying firms. The mean disclosure score ranges from 6.1 to 7.4 forward-looking earnings sentences for firms in the high-disclosure category; whilst the mean disclosure score ranges from 1.4 to 1.8 forward-looking earnings sentences for firms in the low-disclosure category. The differences in means between high- and low-disclosure scores are statistically significant at the 1% level (not reported in Table 1).

Panels D and E of Table 1 show that the median disclosure scores for the high- and low-disclosure groups is quite similar across high- and low-growth firms. The median disclosure score is six forward-looking earnings sentences for firms in the high-disclosure category; whilst it is two forward-looking earnings sentences for firms in the low-disclosure category. In addition, the differences in means between high- and low-disclosure scores are statistically significant at the 1% level (not reported in Table 1). Panels D and E also show that the median return in period $t$ for high-growth firms (3.3%) is slightly lower than that for low-growth firms (4.3%). The median current and future earnings change is positive for both high- and low-growth firms. Finally, the median future earnings changes of low-growth firms are higher than those of high-growth firms.

## 6. Main empirical results

Our main empirical results are based on pooled regressions for the sample period 1996–2002.[13] Table 2 reports these results. Column 2 reports the results for the full sample. Column 3 (4) reports the results for firms in the high- (low-) growth firms. Heteroscedasticity-consistent p-values are given in parentheses.

### *6.1. Disclosure and prices leading earnings*

Table 2, column 2 presents the results of estimating equation (4) for the full sample. As expected, the coefficient associated with $X_t$ is positive and significant. The coefficient for $X_t$ is 0.80 with a p-value of 0.001. However, the coefficient for $X_{t3}$ is –0.06 with a p-value of 0.767. This suggests that current stock price is positively associated with current earnings changes, but not with future earnings change. So there is no evidence that prices lead earnings for low-disclosure firms that pay no dividends.

The incremental predictive value of high forward-looking earnings disclosures for anticipating future earnings is given by the coefficient on

[12] The percentages are calculated as 56% = 1655/2952 and 32% = 174/551.

[13] Hanlon et al. (2007: 16) argue that 'future earnings response regressions are likely to suffer from both cross-sectional correlation (correlation across firms within a year) and time series correlation (over time within the same firm)'. We follow the method recommended by Petersen (2008) and used in Hanlon et al. (2007) by including year dummies to control for the time series correlation and by allowing for error clustering within firms (Rogers standard errors) to control for the cross-sectional correlation.

**Table 2**
**Regression results: the effect of disclosure and dividend propensity on prices leading earnings**

| *Independent variable* (1) | *Full sample* (2) | *High-growth firms* (3) | *Low-growth firms* (4) |
|---|---|---|---|
| Intercept | 0.01 | 0.04 | –0.03 |
| | (0.710) | (0.405) | (0.464) |
| $X_t$ | 0.80*** | 0.42 | 1.12*** |
| | (0.001) | (0.356) | (0.001) |
| $X_{t3}$ | –0.06 | –0.63** | 0.44*** |
| | (0.767) | (0.031) | (0.005) |
| $R_{t3}$ | –0.10*** | –0.12*** | –0.12*** |
| | (0.001) | (0.001) | (0.001) |
| $AG_t$ | 0.07 | 0.06 | 0.09 |
| | (0.158) | (0.325) | (0.353) |
| $EP_{t-1}$ | 0.49*** | 0.12 | 0.83*** |
| | (0.009) | (0.737) | (0.001) |
| $D$ | –0.09* | –0.18** | –0.01 |
| | (0.086) | (0.011) | (0.874) |
| $D*X_t$ | 0.62* | 1.78*** | 0.21 |
| | (0.066) | (0.005) | (0.623) |
| $D*X_{t3}$ | 0.68** | 1.63** | 0.19 |
| | (0.031) | (0.036) | (0.467) |
| $D*R_{t3}$ | 0.06 | 0.10 | 0.04 |
| | (0.175) | (0.116) | (0.494) |
| $D*AG_t$ | 0.06 | 0.11 | 0.05 |
| | (0.483) | (0.382) | (0.722) |
| $D*EP_{t-1}$ | 0.19 | 0.69 | 0.16 |
| | (0.488) | (0.166) | (0.508) |
| $Div$ | 0.05 | 0.07 | 0.06 |
| | (0.241) | (0.259) | (0.203) |
| $Div*X_t$ | 0.92*** | 2.01** | 0.48 |
| | (0.009) | (0.012) | (0.194) |
| $Div*X_{t3}$ | 0.51** | 1.62*** | –0.13 |
| | (0.025) | (0.001) | (0.502) |
| $Div*R_{t3}$ | 0.06* | 0.07 | 0.09* |
| | (0.066) | (0.171) | (0.058) |
| $Div*AG_t$ | 0.07 | 0.06 | 0.06 |
| | (0.296) | (0.455) | (0.562) |
| $Div*EP_{t-1}$ | 0.21 | 0.34 | –0.04 |
| | (0.389) | (0.480) | (0.886) |
| $D*Div$ | 0.10 | 0.13 | 0.04 |
| | (0.127) | (0.136) | (0.674) |
| $D*Div*X_t$ | –0.50 | –2.19** | –0.03 |
| | (0.290) | (0.039) | (0.961) |
| $D*Div*X_{t3}$ | –0.44 | –1.68* | 0.07 |
| | (0.220) | (0.064) | (0.845) |

**Table 2**
**Regression results: the effect of disclosure and dividend propensity on prices leading earnings** *(continued)*

| *Independent variable* (1) | *Full sample* (2) | *High-growth firms* (3) | *Low-growth firms* (4) |
|---|---|---|---|
| $D*Div*R_{t3}$ | –0.05 | –0.08 | –0.03 |
| | (0.357) | (0.293) | (0.664) |
| $D*Div*AG_t$ | –0.11 | –0.10 | –0.13 |
| | (0.288) | (0.461) | (0.397) |
| $D*Div*EP_{t-1}$ | –0.48 | –0.66 | –0.56* |
| | (0.158) | (0.316) | (0.098) |
| Observations | 3,503 | 1,770 | 1,733 |
| $R^2$ | 0.157 | 0.176 | 0.183 |

Table 2 reports regression results. The results for firms in high- and low-growth firms are reported in columns (3) and (4), respectively. The 'Full Sample' results in column (2) combine both types of firms in a single panel. P-values are reported in parentheses. Returns, $R_t$, is calculated as buy-and-hold returns from eight months before the financial year-end to four months after the financial year-end. $R_{t3}$ is the aggregated three years' future returns. The earnings variable, $X_t$, is defined as earnings change per share deflated by the share price four months after the end of the financial year $t$–1. $X_{t3}$ is the aggregated three years' future earnings change. Earnings measure is the *Worldscope* item 01250 which is operating income before all exceptional items. $EP_{t-1}$ is defined as period $t$–1's earnings over price four months after the financial year-end of period $t$–1. $AG_t$ is the growth rate of total book value of assets for period $t$ (*Datastream* item 392). Firms with a disclosure score in the top (bottom) 50% of the distribution of disclosure scores are defined as high- (low-) disclosure firms. The dummy variable, $D$, is set equal to 1 (0) for high- (low-) disclosure firms. Dividends measure is dividends per share, *Worldscope* item 05101. The dummy variable, *Div*, is set equal to 1 (0) for firms paying (not paying) dividends at the current year. High-growth firms are defined as those having below median levels of BTMV, while low-growth firms are defined as those having above median levels of BTMV. *BTMV* ratio is calculated as the inverse of the market-to-book value of equity ratio (*Datastream* item *MTBV*). The significance levels (two-tail test) are: * = 10%, ** = 5%, *** = 1%.

$D*X_{t3}$. The coefficient on $D*X_{t3}$ is 0.68 with a p-value of 0.031. The significantly positive coefficient suggests that high-disclosure firms that pay no dividends exhibit higher levels of share price anticipation of earnings three years ahead than low-disclosure firms that pay no dividends. Thus the effect of disclosure on prices leading earnings is in line with the prior findings of Lundholm and Myers (2002) and Gelb and Zarowin (2002). Based on these findings, we accept hypothesis 1.

Table 2, columns 3 and 4 reveal a number of significant differences between high- and low-growth firms. The current earnings variable exhibits a higher ERC for low-growth firms than high-growth firms. The coefficient on $X_t$ is 1.12 with a p-value of 0.001 for low-growth firms, while it is positive (0.42), but insignificant, for high-growth firms. The difference between the two coefficients is statistically significant at the 1% level.

We find no evidence of share price anticipation of earnings for high-growth firms that pay no dividends and that provide low voluntary disclosure. For these firms, we obtain a statistically significant *negative* coefficient on $X_{t3}$. This indicates that the market is unable to anticipate future earnings changes for high-growth firms that provide low voluntary disclosure in their annual report discussion sections and that do not pay dividends in the current year. In contrast there is strong evidence that low-growth low-disclosure firms that do not pay dividends do exhibit share price anticipation of earnings for three years ahead. The coefficient on $X_{t3}$ is positive and significant at the 1% level.

Looking at the effect of disclosure on prices leading earnings, we find that the coefficient on $D*X_{t3}$ for high-growth firms is 1.63 with a p-value of 0.036. This coefficient indicates that narrative forward-looking earnings disclosures in high-growth firms' annual reports improve the market's ability to anticipate future earnings change three years ahead. Based on this result we accept hypothesis 2.

In contrast, there appears to be no significant effect of disclosure on share price anticipation of earnings for low-growth firms that do not pay dividends. The coefficient on $D*X_{t3}$, for high-disclosure low-growth firms is 0.19 with a p-value of 0.467. Thus, the voluntary disclosures of low-growth firms do not appear to improve the stock market's ability to anticipate future earnings changes. This leads us to reject hypothesis 3.

Overall our evidence for low-growth firms suggests that the market is able to forecast future earnings changes, but this ability is neither linked to nor improved by forward-looking earnings statements in annual report narrative sections. The evidence for high-growth firms supports the view that the market has particular difficulties in forecasting firms' future earnings changes. However, this difficulty is partially overcome by increasing the number of forward-looking earnings information in annual report narrative sections.

We also test for differences between high-growth firms and low-growth firms. We test the extent to which the association between share price anticipation of earnings and disclosure is significantly stronger for high-growth firms than for low-growth firms. We perform this test by including all high- and low-growth firms in one dataset. Then, before running our analyses, we create a dummy variable to equal 1 for high-growth firms and zero otherwise. Finally, we interact the high-growth dummy variable throughout the model. We find a positive and significant coefficient on $Growth{*}D{*}X_{t3}$ of 0.50 with a p-value of 0.020 (not reported in Table 2). This suggests that that the strength of the degree of association between share price anticipation of earnings and voluntary disclosure is stronger for high-growth firms than for low-growth firms. This leads us to accept hypothesis 4.

*6.2. Dividends and prices leading earnings*

As discussed earlier, voluntary disclosure is one of the ways that firms provide information about future earnings to the market. Another way that firms can provide information is through dividend policy. Here we examine the extent to which dividend propensity improves the stock market's ability to anticipate future earnings changes. The incremental predictive effect of dividend propensity on share price anticipation of earnings is given by the coefficient on $Div{*}X_{t3}$.

Column 2, Table 2 shows that the coefficient on $Div{*}X_{t3}$ is positive and significant. This coefficient is 0.51 with a p-value of 0.025. This indicates that paying dividends improves the market's ability to anticipate future earning changes. The significant positive coefficient indicates that low-disclosure firms that pay dividends exhibit higher levels of share price anticipation of earnings than low-disclosure firms that pay no dividends. The significant effect of dividend propensity on prices leading earnings is in line with the findings of Hanlon et al. (2007). Based on these findings, we accept hypothesis 5.

Table 2, column 3 shows that the coefficient on $Div{*}X_{t3}$ for high-growth firms is 1.62 with a p-value of 0.001. This coefficient demonstrates that high-growth dividend-paying firms exhibit significantly higher levels of share price anticipation of earnings than high-growth non-dividend-paying firms. Based on this result we accept hypothesis 6.

In contrast, Table 2, column 4 shows that the coefficient on $Div{*}X_{t3}$ for low-growth firms is negative and insignificant. Thus, there is no evidence that the dividend propensity of low-growth firms is associated with an improvement in the market's ability to anticipate future earnings changes. Based on this result we reject hypothesis 7.

Overall our evidence indicates that there is a material difference between high- and low-growth firms in the association between dividend propensity and share price anticipation of earnings. For high-growth firms we find that the ability of the market to anticipate future earnings changes is significantly greater when the firm pays dividends in the current year. We find no such evidence for low-growth firms.

We test for a difference between high-growth and low-growth firms by interacting a dummy variable for high growth throughout the model. We find a positive and significant coefficient on $Growth{*}Div{*}X_{t3}$ of 1.41 with a p-value of 0.001 (not reported in Table 2). This suggests that the strength of the degree of association between share price anticipation of earnings and dividend propensity is stronger for high-growth firms than for low-growth firms. This leads us to accept hypothesis 8.

Summarising the results for the above hypotheses (H1–H8), we find that both dividend propensity and high disclosure are positively associated with share price anticipation of earnings for high-growth firms. However, there is no evidence of similar effects for low-growth firms.

*6.3. The joint effect of disclosure and dividends on prices leading earnings*

We now turn to the hypotheses which are concerned with the joint effects of high disclosure and dividend propensity on prices leading earnings. The incremental predictive value of both high disclosure and dividend propensity for anticipating future earnings is given by the coefficient on $D{*}Div{*}X_{t3}$. Column 2, Table 2 shows some evidence of a substitution effect. The coefficient on the interaction variable $D{*}Div{*}X_{t3}$ is negative (–0.44) and statistically insignificant at an accepted level. This indicates that both disclosure and dividends provide related information, but some of the information is common to both, i.e. the effects are 'partially additive'. Our best estimate for the combined effect of disclosure and dividend is 0.75 (i.e. 0.68 + 0.51– 0.44] which is smaller than the sum of the coefficients on $D{*}X_{t3}$ and $Div{*}X_{t3}$ [1.19 = 0.68 + 0.51] with a p-value of 0.001 (not reported in Table 2). However, because of the insignificant coefficient

on $D*Div*X_{t3}$, it is not safe to assume that the combined effect of disclosure and dividends is additive. Therefore, we reject hypothesis 9.

For high-growth firms, we also find some evidence of a substitution effect. The coefficient on the interaction variable $D*Div*X_{t3}$ is negative (–1.68) and significantly significant at the 10% level. This indicates that disclosure and dividends provide related information, but some of the information is common to both. Thus, it is not safe to assume that the combined effect of high disclosure and dividend propensity is perfectly additive. Our best estimate is that the combined effect of dividend propensity and high disclosure is 1.57 (i.e. 1.63 + 1.62 – 1.68) with a p-value of 0.001 (not reported in Table 2), which is below the first order effect for high disclosure and below the first order effect of dividend propensity. In this case, the inference is that both dividend propensity and voluntary disclosure are strict substitutes for high-growth firms. Based on these findings we reject hypothesis 10.

As the results for the low-growth firms indicate that there is no first order effects either for dividend propensity or voluntary disclosure, we find that the coefficient on $D*Div*X_{t3}$ is positive (close to zero) and statistically insignificant at an accepted level. Based on these findings we reject hypothesis 11.

Finally, we test for differences between high-growth and low-growth firms by including an additional dummy variable for high-growth firms. We test the extent to which the association between the joint effect of voluntary disclosure and the payments of dividends on prices leading earnings are significantly stronger for high-growth firms than for low-growth firms. The analysis shows a negative significant coefficient on $Growth*D*Div*X_{t3}$ of –1.83 with a p-value of 0.067 (not reported in Table 2). This significantly negative coefficient indicates that dividend propensity and high voluntary disclosure are strict substitutes for high-growth firms. We do not find such evidence for low-growth firms. This leads us to reject hypothesis 12.

## 7. Robustness analysis

In this section we examine the sensitivity of our results to the determinants of the earnings response coefficients. Lundholm and Myers (2002) examine a number of determinants of *current* earnings response coefficient when exploring the association between share price anticipation of earnings and corporate disclosure. These determinants include loss status, growth, beta, earnings persistence, size and the sign of the current return. Hanlon et al. (2007) examine a similar set of determinants when exploring the association between share price anticipation of earnings and dividend propensity. The results of both studies (after the inclusion of control variables) remain consistent with the original findings indicating that these control variables do not drive the association between disclosure (and dividends) and share price anticipation of earnings.

Schleicher et al. (2007) provide evidence that the association between annual report narratives and share price anticipation of earnings is not the same for profitable and unprofitable firms. They find that the ability of stock returns to anticipate the next year's earnings change is significantly stronger for high-disclosure unprofitable firms. They do not find the same result for profitable firms. Therefore, based on the results in Schleicher et al. (2007), we examine the sensitivity of our results to firm profitability status. Similar to Schleicher et al. (2007) we define a loss (profit) as negative (positive) operating income before all (operating and non-operating) exceptional items (*Worldscope* item 01250).

To examine the effect of losses on the association between disclosure, dividends and prices leading earnings, we divide our sample into two categories; unprofitable firms and profitable firms. Then, we run our regression model (equation 4) for each category. The results are reported in Table 3.

Consistent with Schleicher et al. (2007), columns 2.1 and 3.1 of Table 3 show that high disclosure increases the market's ability to anticipate future earnings changes for unprofitable firms – but not for profitable firms. The coefficient on $D*X_{t3}$ is positive (1.01) and statistically significant at the 5% level for unprofitable firms, whilst it is smaller and insignificant for profitable firms.

Table 3 shows that several of our previous findings still hold after separating our sample into unprofitable firms and profitable firms. In particular, column 3.3 shows that the market is able to anticipate future earnings changes for low-growth profitable firms (the coefficient on $X_{t3}$ is significantly positive at the 5% level). This ability is neither linked to nor improved by high disclosure or dividend propensity. In addition, columns 2.2 and 3.2 show that the market has particular difficulties in anticipating future earnings changes for high-growth profitable and unprofitable firms. The coefficient on $X_{t3}$ for these firms is negative.

Consistent with results in Table 2, we find that the effect of high disclosure on prices leading earnings is positive for high-growth firms regardless of their profitability status. The coefficient on $DX_{t3}$ is 0.93 with a p-value of 0.209 for high-growth unprofitable firms and 1.47 with a p-value of 0.160 for high-growth profitable firms. Table 3 also shows that the effect of disclosure on prices leading earnings for high-growth firms is greater than the effect for low-growth firms regardless of the profitability of the firms. In particular, the

**Table 3**
**Robustness analysis: comparing unprofitable with profitable firms**

| *Independent variable (1)* | *Unprofitable firms (2)* | | | *Profitable firms (3)* | | |
|---|---|---|---|---|---|---|
| | *Full sample* (2.1) | *High-growth firms* (2.2) | *Low-growth firms* (2.3) | *Full sample* (3.1) | *High-growth firms* (3.2) | *Low-growth firms* (3.3) |
| Intercept | –0.14* (0.087) | –0.01 (0.957) | –0.26*** (0.001) | 0.08 (0.285) | 0.06 (0.556) | 0.05 (0.617) |
| $X_t$ | –0.28 (0.343) | –0.90 (0.122) | 0.21 (0.470) | 1.13*** (0.009) | 1.50** (0.029) | 1.36** (0.014) |
| $X_{t3}$ | –0.56** (0.025) | –1.12** (0.019) | 0.10 (0.645) | 0.18 (0.585) | –0.28 (0.451) | 0.68** (0.015) |
| $R_{t3}$ | –0.09*** (0.006) | –0.12** (0.012) | –0.07* (0.051) | –0.10** (0.028) | –0.06 (0.316) | –0.19*** (0.001) |
| $AG_t$ | 0.09 (0.141) | 0.05 (0.522) | 0.13 (0.207) | 0.08 (0.440) | 0.11 (0.280) | 0.03 (0.885) |
| $EP_{t-1}$ | –0.30 (0.269) | –0.88 (0.108) | 0.18 (0.468) | 0.51 (0.217) | 0.35 (0.657) | 0.89* (0.064) |
| $D$ | –0.13 (0.048) | –0.16* (0.098) | 0.02 (0.864) | –0.14 (0.177) | –0.19 (0.224) | 0.07 (0.654) |
| $D*X_t$ | 0.69 (0.125) | 2.24** (0.030) | 0.29 (0.512) | 0.74 (0.270) | 0.24 (0.802) | 0.60 (0.481) |
| $D*X_{t3}$ | 1.01** (0.011) | 0.93 (0.209) | 0.59 (0.101) | 0.58 (0.184) | 1.47 (0.160) | –0.08 (0.809) |
| $D*R_{t3}$ | 0.10 (0.109) | 0.10 (0.242) | 0.09 (0.178) | 0.01 (0.965) | 0.02 (0.849) | –0.02 (0.873) |
| $D*AG_t$ | 0.02 (0.854) | 0.03 (0.763) | –0.03 (0.808) | 0.28 (0.118) | 0.23 (0.358) | 0.33 (0.232) |
| $D*EP_{t-1}$ | 0.29 (0.488) | 0.66 (0.325) | 0.60 (0.134) | 0.38 (0.429) | 0.98 0.3029 | –0.42 (0.421) |
| $Div$ | 0.06 (0.494) | –0.04 (0.804) | 0.18* (0.084) | –0.04 (0.580) | –0.01 (0.978) | –0.04 (0.706) |
| $Div*X_t$ | 0.85* (0.075) | 1.03 (0.275) | 1.05** (0.036) | 0.90* (0.098) | 2.15 (0.023) | 0.43 (0.511) |
| $Div*X_{t3}$ | 0.53 (0.248) | 1.08 (0.390) | 0.08 (0.883) | 0.32 (0.370) | 1.51 (0.004) | –0.37 (0.219) |
| $Div*R_{t3}$ | 0.24** (0.016) | 0.33** (0.044) | 0.20** (0.039) | 0.04 (0.402) | –0.02 (0.775) | 0.14** (0.029) |
| $Div*AG_t$ | –0.01 (0.974) | –0.19 (0.692) | 0.40 (0.149) | 0.07 (0.510) | 0.01 (0.935) | 0.11 (0.604) |
| $Div*EP_{t-1}$ | 0.19 (0.674) | –0.13 (0.894) | 0.79 (0.229) | 0.38 (0.395) | 0.55 (0.528) | 0.03 (0.958) |
| $D*Div$ | 0.05 (0.661) | 0.23 (0.357) | –0.25* (0.088) | 0.17 (0.118) | 0.19 (0.250) | –0.01 (0.930) |
| $D*Div*X_t$ | –2.02* (0.098) | –1.81 (0.583) | –3.91*** (0.001) | –0.90 (0.254) | –1.90 (0.136) | –0.55 (0.566) |

**Table 3**
**Robustness analysis: comparing unprofitable with profitable firms** *(continued)*

| *Independent variable (1)* | *Unprofitable firms (2)* | | | *Profitable firms (3)* | | |
|---|---|---|---|---|---|---|
| | *Full sample* (2.1) | *High-growth firms* (2.2) | *Low-growth firms* (2.3) | *Full sample* (3.1) | *High-growth firms* (3.2) | *Low-growth firms* (3.3) |
| $D{*}Div{*}X_{t3}$ | –0.23 | 1.68 | –1.00 | –0.40 | –1.79 | 0.36 |
| | (0.749) | (0.332) | (0.184) | (0.404) | (0.113) | (0.384) |
| $D{*}Div{*}R_{t3}$ | –0.23* | –0.40** | –0.26** | 0.02 | 0.03 | 0.03 |
| | (0.052) | (0.037) | (0.023) | (0.765) | (0.761) | (0.794) |
| $D{*}Div{*}AG_t$ | 0.47 | 0.86 | –0.41 | –0.34* | –0.23 | –0.42 |
| | (0.251) | (0.173) | (0.227) | (0.068) | (0.367) | (0.144) |
| $D{*}Div{*}EP_{t-1}$ | –2.15* | –1.96 | –4.66*** | –0.88 | –1.42 | –0.14 |
| | (0.072) | (0.592) | (0.001) | (0.096) | (0.186) | (0.814) |
| Observations | 439 | 232 | 207 | 3,064 | 1,538 | 1,526 |
| $R^2$ | 0.170 | 0.304 | 0.164 | 0.143 | 0.162 | 0.174 |

Table 3 reports robustness results. The results for unprofitable and profitable firms are reported in columns (2) and (3), respectively. In each category, we report the results for the full sample, high-growth firms and low-growth firms. P-values are reported in parentheses. Returns, $R_t$, is calculated as buy-and-hold returns from eight months before the financial year-end to four months after the financial year-end. $R_{t3}$ is the aggregated three years' future returns. The earnings variable, $X_t$, is defined as earnings change per share deflated by the share price four months after the end of the financial year $t$–1. $X_{t3}$ is the aggregated three years' future earnings change. Earnings measure is the *Worldscope* item 01250 which is operating income before all exceptional items. $EP_{t-1}$ is defined as period $t$–1's earnings over price four months after the financial year-end of period $t$–1. $AG_t$ is the growth rate of total book value of assets for period $t$ (*Datastream* item 392). Firms with a disclosure score in the top (bottom) 50% of the distribution of disclosure scores are defined as high- (low-) disclosure firms. The dummy variable, $D$, is set equal to 1 (0) for high- (low-) disclosure firms. Dividends measure is dividends per share, *Worldscope* item 05101. The dummy variable, $Div$, is set equal to 1 (0) for firms paying (not paying) dividends at the current year. High-growth firms are defined as those having below median levels of BTMV, while low-growth firms are defined as those having above median levels of BTMV. *BTMV* ratio is calculated as the inverse of the market-to-book value of equity ratio (*Datastream* item *MTBV*). The significance levels (two-tail test) are: * = 10%, ** = 5%, *** = 1%.

coefficient on $DX_{t3}$ is higher for high-growth unprofitable firms than for low-growth unprofitable firms. However, the difference between high-growth unprofitable firms and low-growth unprofitable firms is not statistically significant (not reported in Table 3).[14] On the other hand, the effect of disclosure on prices leading earnings for high-growth profitable firms is significantly greater than the effect for low-growth profitable firms (significant at the 10% level; not reported in Table 3). The results suggest that the strength of the degree of association between share price anticipation of earnings and voluntary disclosure is greater for high-growth firms than for low-growth firms.

Consistent with results in Table 2, we find that the effect of dividend propensity on prices leading earnings is positive for high-growth firms, while it is negative or close to zero for low-growth firms. However, the effect for high-growth firms is only statistically significant for high-growth profitable firms at the 1% level.

The effect of dividend propensity on prices leading earnings is significantly greater for high-growth firms than for low-growth firms, in both profitability subsamples. For unprofitable firms the coefficient on $Div{*}X_{t3}$ is 1.08 with a p-value of 0.390 for high-growth unprofitable firms, while it is 0.08 with a p-value of 0.883 for low-growth unprofitable firms. The difference between high and low-growth unprofitable firms is statistically significant at the 10% level (not reported in Table 3). For profitable firms the coefficient on $Div{*}X_{t3}$ is

[14] As mentioned earlier, we statistically test the actual differences between high- and low-growth firms by interacting a dummy variable for high growth throughout Equation (4).

1.51 and significant at the 1% level for high-growth firms, while it is negative and insignificant for low-growth firms. The difference between high and low-growth profitable firms is statistically significant at the 1% level (not reported in Table 3). These results confirm that the strength of the degree of association between share price anticipation of earnings and dividend propensity is greater for high-growth firms than for low-growth firms.

A particularly interesting feature of Table 3 is the way that the joint effect of dividend propensity and disclosure varies between profitable and unprofitable firms. In particular the results suggest some (weak) evidence of a complementary effect for high-growth unprofitable firms. The subsitution effect that we reported in Table 2 is confirmed in the high-growth profitable subsample, although it is no longer statistically significant. The difference between high-growth unprofitable and high-growth profitable firms for the coefficient on $D^*Div^*X_{t3}$ is significant at the 1% level (not reported in Table 3).

Finally, it is worth noting that for high-growth unprofitable firms the combined effects of high disclosure and dividend propensity are very considerable (0.93 + 1.08 + 1.68 = 3.69). High-growth unprofitable firms are the firms for which current earnings is least relevant (revealed by a negative current earnings response coefficient (ERC)), and for which the combined effects of narrative disclosure and dividend propensity are the most useful for predicting future earnings.

## 8. Conclusion

This paper builds on literature that examines the link between narrative disclosures and prices leading earnings.

We extend this work in two important ways. First, it is well known that financial policy choices, such as dividend signalling, potentially offer an alternative set of devices for conveying value-relevant information to the market. In particular firms that pay dividends may use changes in dividends to signal future profitability. We investigate whether firms that pay dividends exhibit higher levels of share price anticipation of earnings, and whether dividend propensity acts as a substitute for narrative disclosure in financial communication.

Second, we investigate whether firm growth characteristics affect the extent to which prices lead earnings, and whether the importance of dividend signalling and narrative disclosures varies between high- and low-growth firms.

Our results show a number of significant differences between high- and low-growth firms. We find that high-growth, low-disclosure firms that pay no dividends exhibit no share price anticipation of earnings. On the other hand, we find that low-growth, low-disclosure firms that pay no dividends exhibit significant share price anticipation of earnings. We also find that dividend propensity and high voluntary disclosure improve the market's ability to anticipate future earnings changes for high-growth firms, but not for low-growth firms.

With regard to the additvity or otherwise of disclosure and dividend propensity we find, for the high-growth firm subsample for which both of the individual effects are significant, that the effects of voluntary disclosure and dividend propensity on the degree of share price anticipation of earnings are not perfectly additive.

This paper is the first to study effect of firm-specific growth characteristics on the association between voluntary disclosure, dividend propensity and prices leading earnings. Whilst we focus only on the growth characteristics of firms it would also be interesting to examine the effect of other firms' characteristics such as risk.

Another interesting issue for future work would be to consider the factors that determine the choice between dividend signalling and increased disclosure as alternative forms of financial communication for high-growth firms. For example, is it the case that firms with potentially high third party disclosure costs are more likely to use dividend signalling?

Finally, it is worth noting that the dividend propensity of UK firms has recently declined (Vieira and Raposo, 2007), although this has to some extent been offset by an increase in share repurchases. Future work could test for a change in the dividend propensity effect following an overall decline in dividend propensity, and it could also test to see if share repurchases are associated with greater share price anticipation of earnings.

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# Non-linear equity valuation

**Ali Ataullah, Huw Rhys and Mark Tippett***

**Abstract**—We incorporate a real option component into the Ohlson (1995) equity valuation model and then use this augmented model to make assessments about the form and nature of the systematic biases that are likely to arise when empirical work is based on linear models of the relationship between the market value of equity and its determining variables. We also demonstrate how one can expand equity valuation models in terms of an infinite series of 'orthogonal' polynomials and thereby determine the relative contribution which the linear and non-linear components of the relationship between equity value and its determining variables make to overall equity value. This procedure shows that non-linearities in equity valuation can be large and significant, particularly for firms with low earnings-to-book ratios or where the undeflated book value of equity is comparatively small. Moreover, it is highly unlikely the simple linear models that characterise this area of accounting research can form the basis of meaningful statistical tests of the relationship between equity value and its determining variables.

**Keywords:** branching process; equity valuation, orthogonal polynomials; real option value; recursion value; scaling

## 1. Introduction

Theoretical developments in capital investment analysis show that real options can play an important role in the capital budgeting process. If a firm has the option of abandoning a poorly performing capital project it can increase the capital project's value well beyond the traditional benchmark given by the present value of its expected future cash flows; likewise, the growth option associated with an unexploited capital project can also be significant when compared with the expected present value of its future cash flows. Moreover, it is now generally accepted that these option values mean that evaluating capital projects exclusively in terms of the present values of their future cash flows can lead to seriously flawed investment decisions – highly profitable capital projects can be overlooked and poorly performing capital projects wrongly implemented. Given this, it is somewhat surprising that both empirical and analytical work on the relationship between equity value and its determining variables continues to be based on models that establish the value of a firm's equity exclusively in terms of the present value of its future operating cash flows and which, therefore, ignore the real option effects associated with the firm's ability to modify or even abandon its existing investment opportunity set. The Ohlson (1995) model, for example, from which much of the empirical work in the area is motivated (Barth and Clinch, 2005: 1) implies that there is a purely linear relationship between the market value of equity and its determining variables. As such, it is based on the implicit assumption that real options are of little significance in the equity valuation process. Given this, it is all but inevitable that when real options do impact on equity values the Ohlson (1995) model will return a systematically biased picture of the relationship between the market value of equity and its determining variables. Fortunately, Ashton et al. (2003) have generalised the Ohlson (1995) model so that it takes account of the real options generally available to firms. Our task here is to use this more general model to determine the likely form and magnitude of the biases that arise under linear equity valuation models like the one formulated by Ohlson (1995).

In the next section we briefly summarise the principal features of the Ashton et al. (2003) equity valuation model and, in particular, some important scale invariance principles on which it is based. Recall here that empirical work in the area is invariably based on market and/or accounting (book) variables that have been normalised or deflated in order to facilitate comparisons between firms of different size. Given this, it is important that one appreciates how these deflation procedures might alter or even distort the underlying levels relationship which exists between the mar-

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ket value of equity and its determining variables. In section 3 we introduce a hitherto unused orthogonal polynomial fitting procedure for identifying the relative contribution which the linear and non-linear components of the relationship between equity value and its determining variables make to overall equity value. The evidence from this procedure is that non-linearities in equity valuation can be large and significant, particularly for firms with low earnings-to-book ratios or where the undeflated book value of equity is comparatively small. In Section 4 we examine the specific nature of the statistical biases which arise when the simple linear models that characterise this area of accounting research are used to model the complex non-linear relationships that actually exist. Our conclusion is that it is highly unlikely these simple linear models can form the basis of meaningful statistical tests of the relationship between equity value and its determining variables. Section 5 contains our summary and conclusions.

## 2. Real options and equity value

The Ashton et al. (2003) model is based on the assumption that the market value of a firm's equity has two components. The first is called the recursion value of equity and is the present value of the future cash flows the firm expects to earn given that its existing investment opportunity set is maintained indefinitely into the future. The Ohlson (1995) equity valuation model is exclusively based on this component of equity value. There is, however, a second component of equity value which the Ohlson (1995) model ignores; namely, the real option (or adaptation) value of equity. This is the option value that arises from a firm's ability to change its existing investment opportunity set by (for example) fundamentally changing the nature of its operating activities (Burgstahler and Dichev, 1997 :188). Ashton et al. (2003) employ this distinction to develop a quasi-supply side generalisation of the Ohlson (1995) model under which the recursion value of equity, $\eta(t)$, evolves in terms of a continuous time branching process (Feller, 1951: 235–237; Cox and Ross, 1976: 149):

$$d\eta(t) = i\eta(t)dt + \sqrt{\eta(t)}dq(t) \tag{1}$$

where i is the cost of equity capital and dq(t) is a Wiener process with a variance parameter of $\zeta$. It then follows that in expectations the proportionate rate of growth in the recursion value of equity will be equal to the cost of equity, or $E_t[d\eta(t)] = i\eta(t)dt$ where $E_t(\cdot)$ is the expectations operator taken at time t. Furthermore, the variance of instantaneous increments in the recursion value of equity will be $\mathrm{Var}_t[d\eta(t)] = \zeta\eta(t)dt$. Moreover, since the literature normally describes the firm's investment opportunity set in terms of a first order autoregressive system of information variables it also follows that the recursion value of equity, $\eta(t)$, will be a linear sum of more primitive variables. A good example of this is provided by the Ohlson (1995: 667–669) linear information dynamics which implies that the recursion value of equity evolves in terms of the book value of equity, b(t), book or accounting earnings, x(t), and an information variable, v(t), in accordance with the following formula:

$$\eta(t) = c_1 b(t) + c_2 x(t) + c_3 v(t) \tag{2}$$

where $c_1$, $c_2$ and $c_3$ are the relevant valuation coefficients.[1]

Ashton et al. (2003) employ the quasi-supply side model summarised above in conjunction with standard no arbitrage conditions and thereby show that the market value of the firm's equity, $P(\eta,B,\theta)$, will have to be:[2]

$$P(\eta,B,\theta) = \eta + \frac{B}{2}\int_{-1}^{1} \exp\left(\frac{-2\theta\eta}{1+z}\right)dz \tag{3}$$

Here B > 0 denotes the value of the firm's adaptation options when the recursion value of equity

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[1] There is nothing particularly unique about the Ohlson (1995) linear information dynamics, despite its widespread use in market-based accounting research (Barth and Clinch, 2005: 1). If, for example, one lets earnings and the information variables evolve in terms of a second order system of linear difference equations it is not hard to show that the recursion value of equity will hinge on both the levels of the earnings and information variables as well as the momentum (or first differences) in both these variables. Moreover, if earnings and the information variable evolve in terms of a third order system of linear difference equations, the recursion value of equity will hinge on the levels, momentum and acceleration (or second differences) in both these variables. One could generalise these results to even higher order systems of difference equations. It suffices here to note that these higher order systems provide an analytical basis for the growing volume of empirical evidence which appears to show that the momentum of variables comprising the firm's investment opportunity set can have a significant impact on the value of equity (Chordia and Shivakumar, 2006).

[2] The Ashton et al. (2003) model assumes the firm does not pay dividends and practises 'clean surplus' accounting. However, Ashton et al. (2004) have generalised the Ashton et al. (2003) model so as to relax both these assumptions. They show that the relationship between equity value and recursion value for a dividend paying/dirty surplus firm has exactly the same convex structure and properties as the relationship between equity value and recursion value for a non-dividend paying/clean surplus firm. However, the mathematics of a dividend paying/dirty surplus firm is much more complicated – generally involving numerical procedures as compared to the closed form solutions available under the Ashton et al. (2003) model. Hence, given the pedagogical disadvantages associated with these numerical procedures and the solutions they entail and the equivalent convex structure and properties of the two models, we develop our subsequent analysis in terms of the non-dividend paying/clean surplus Ashton et al. (2003) firm without any loss in generality.

falls away to nothing and

$$\theta = \frac{2i}{\mathrm{Var}_t[dq(t)]} = \frac{2i}{\zeta}$$

is a risk parameter that captures the relative stability with which the recursion value of equity evolves over time.[3] Note that $\eta$, the first term on the right-hand side of this equation, is the recursion value of equity on which the Ohlson (1995) equity valuation model is exclusively based. We have previously noted, however, that firms invariably have the option of changing their investment opportunity sets and this gives rise to the real option component of equity value captured by the integral term in the above valuation formula. Here it is important to note that as the variability ($\zeta$) of the recursion value increases relative to the cost of equity (i), the term $\exp(\frac{-2\theta\eta}{1+z})$ grows in magnitude and the real option value of equity becomes larger as a consequence. Similarly, the real option value of equity falls as the variability of the recursion value declines relative to the cost of equity. In other words, when the rate of growth in recursion value clusters closely around the cost of equity it is unlikely the catastrophic events which will induce the firm to exercise its real options will arise. In these circumstances, the small probability of these options ever being exercised will mean that the real option value of equity will also have to be comparatively small.

Here we need to note, however, that empirical work in the area is invariably based on market and/or accounting (book) variables that have been normalised or deflated in order to facilitate comparisons between firms of different size. Given this, suppose one defines the normalised recursion value, $h(t) = \frac{\eta(t)}{w}$, where w is the normalising factor. It then follows that increments in the normalised recursion value will evolve in terms of the process:

$$dh(t) = ih(t)dt + \sqrt{h(t)}dv(t) \tag{4}$$

where

$$dv(t) = \frac{1}{\sqrt{w}}.dq(t)$$

is a Wiener process with variance parameter $\frac{\zeta}{w}$. Note that increments in the normalised recursion value, h(t), will have a mean and variance of $E_t[dh(t)] = ih(t)dt$ and $\mathrm{Var}_t[dh(t)] = h(t)\mathrm{Var}_t[dv(t)] = h(t)\frac{\zeta}{w}dt$, respectively. This in turn implies that the risk parameter for the normalised recursion value will be

$$\frac{2i}{\mathrm{Var}_t[dv(t)]} = \frac{2iw}{\zeta} = \theta w.$$

Furthermore, if one works in terms of this normalised variable, h(t), rather than in levels, $\eta(t)$, it follows that the value of equity will have to be:

$$P(h,\frac{B}{w},\theta w) = h + \frac{B}{2w}\int_{-1}^{1}\exp(\frac{-2\theta wh}{1+z})dz \tag{5}$$

Here it is important to note that $\eta(t) = wh(t)$ and so, the normalised value of equity will satisfy the following important property:

$$wP(h,\frac{B}{w},\theta w) = \eta + \frac{B}{2}\int_{-1}^{1}\exp(\frac{-2\theta\eta}{1+z})dz = P(\eta,B,\theta) \tag{6}$$

Formally, this result means that the Ashton et al. (2003) equity valuation model is scale-invariant under all dilations, w (Borgnat et al., 2002: 181).[4]

Now consider a firm for which all variables have been deflated by the book value of equity, w = B, as at some *fixed* date or that $h(t) = \frac{\eta(t)}{B}$ in the scaled version of the Ashton et al. (2003) model given earlier. Moreover, assume, for purposes of illustration, that the firm's adaptation options involve selling off its existing investment opportunities at their book values as recorded on the balance sheet on this fixed date and using the proceeds to move into alternative lines of business.[5] These consider-

[3] The Ashton et al. (2003) model follows previous work (Burgstahler and Dichev, 1997: 190–195) in assuming that the adaptation value of equity, B, is a given known constant. There are, however, a variety of ways in which one could allow B to vary stochastically in the Ashton et al. (2003) valuation formula given here. The most parsimonious of these lets B evolve as a random walk – although even here one could have an error structure that allows for the variance of increments in adaptation value to hinge on the current adaptation value itself, thereby preventing adaptation value from becoming negative. Standard no-arbitrage conditions will then lead to exactly the same Hamilton-Jacobi-Bellman equation from which the Ashton et al. (2003) valuation formula given above is derived.

[4] If, for all real x, $\lambda$ and some $\Delta$, a function $f(\cdot)$, satisfies the property $\lambda^{\Delta}f(x) = f(\lambda x)$ then it is said to be scale invariant for all dilations, $\lambda$, with a scaling dimension of $\Delta$ (Borgnat, et al., 2002: 181). See Alexander and Nogueira (2006: 4–6) for examples of option models that violate the scale-invariance property given here.

[5] This assumption is widely used and accepted in the empirical work of the area. Burgstahler and Dichev (1997: 195), for example, define the book value of equity at the beginning of the interval covered by the profit and loss statement 'as the measure of adaptation value for the year'. Barth et al. (1998: 1–2), Collins et al. (1999: 32), Ashton et al. (2003: 427) and Cotter and Donnelly (2006: 11) amongst many others also invoke this assumption. All these authors base their argument for using opening book value as the adaptation value for the year on the fact that the impairment testing procedures summarised in accounting standards like FRS 10: *Goodwill and Intangible Assets*, FRS 11: *Impairment of Fixed Assets and Goodwill*, FRS 15: *Tangible Fixed Assets* and the Fourth Schedule of the Companies Act (or their North American equivalents) require assets appearing on a firm's balance sheet to be carried at no more than their recoverable amounts. This in turn means it is highly likely the book value of equity at some given fixed point in time will constitute a reasonable approximation for the adaptation value of equity over short intervals surrounding that fixed point in time. See Burgstahler and Dichev (1997: 187–194) for an elaboration of these arguments.

ations will mean the market value of equity (per unit of book value) evolves in terms of the following formula:

$$\frac{P(\eta,B,\theta,)}{B} = h + \frac{1}{2}\int_{-1}^{1} \exp(\frac{-2\theta Bh}{1+z})dz = P(h,1,\theta B) \quad (7)$$

where $\frac{P(\eta,B,\theta)}{B}$ is the ratio of the market value of equity to its book value on the given fixed date. This result implies that there will be a convex relationship between the market-to-book ratio of equity, P(h,1,θB), and the recursion value of equity divided by the book value of equity, h.[6] Suppose that one assumes the existence of a linear relationship between the market value of equity and its determining variables, as is the case with the Ohlson (1995) model. These linear models are based on the implicit assumption that the real options generally available to firms have no role to play in the equity valuation process. However, if one uses a linear model to approximate the relationship between equity value and its determining variables when, in fact, real options do impact on equity values then it is all but inevitable there will be systematic differences between the actual market values and those predicted by the linear model. One can illustrate the point being made here by approximating the market-to-book ratio given in the scaled version of the Ashton et al. (2003) model by a linear function over its entire domain. Our approximating procedures make use of an inner product (Hilbert) space using the Laguerre polynomials as a basis and are summarised in further detail in the Appendix. These procedures show that the best linear approximation to the equity valuation function [P(h,1,θB)] over the semi-infinite real line will be:

$$P(h,1,\theta B) = h + \frac{1}{2}\int_{-1}^{1} \exp(\frac{-2\theta Bh}{1+z})dz \approx \frac{1}{1+\theta B} \quad (8)$$

$$+ [1 + \frac{\theta B}{1+\theta B} + \theta B \log(\frac{\theta B}{1+\theta B})].h$$

As a particular example, consider a firm whose normalised risk parameter is $\theta B = 2$ in which case substitution shows that the best linear approximation to the equity valuation function will be:[7]

$$P(h,1,2) = h + \frac{1}{2}\int_{-1}^{1} \exp(\frac{-4h}{1+z})dz \approx 0.3333 + 0.8557h \quad (9)$$

Figure 1 contains a diagrammatic summary of these results. The upward sloping line emanating from the origin at a 45 degree angle is the normalised recursion value of equity, $h = \frac{\eta}{B}$. The downward sloping curve which asymptotes towards the recursion value axis is the normalised real option value of equity,

$$\frac{1}{2}\int_{-1}^{1} \exp(\frac{-4h}{1+z})dz.$$

The sum of the normalised recursion and normalised real option values is the market value of equity per unit of book value,

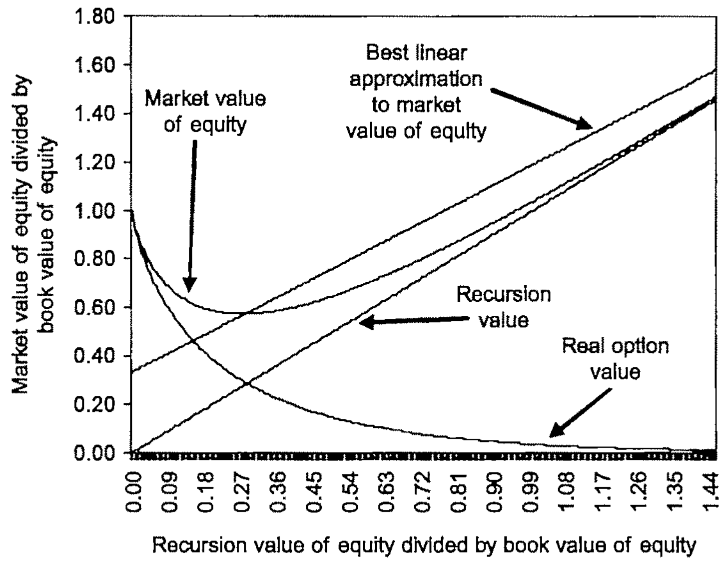
$$P(h,1,2) = h + \frac{1}{2}\int_{-1}^{1} \exp(\frac{-4h}{1+z})dz,$$

and is represented by the convex curve which asymptotes towards the 45 degree line representing the normalised recursion value of equity. Here it is important to observe that as the normalised recursion value increases in magnitude, the market value of equity (per unit of book value) at first declines before reaching a minimum and then gradually increases in magnitude. This arises because at small recursion values the decline in real option value will be much larger than the increase in the recursion value itself. This, in turn, means that the best linear approximation, 0.3333 + 0.8557h, to the overall market value of equity will bear a particular relationship to the market value of equity, P(h,1,2). Hence, when the normalised recursion value of equity is low ($h < 0.30$) then the difference between the linear approximation and the market value of equity, (0.3333 + 0.8557h) – P(h,1,2), is negative. Beyond this point (that is, $h > 0.30$) the difference between the linear approximation and the market value of equity becomes positive before falling away again and becoming negative when

[6] The early evidence of Burgstahler and Dichev (1997: 199–205) and Burgstahler (1998: 339) shows that at an aggregate level there is a highly non-linear and generally convex relationship between the ratio of the market value of equity to the book value of equity and the ratio of earnings attributable to equity to the book value of equity for US equity securities over the 20-year period ending in 1994. Likewise, Ashton et al. (2003: 429–430) show that an almost identical convex relationship exists at an aggregate level between the market value of equity and the earnings attributable to equity for UK equity securities over the period from 1987 to 1998. Finally, Di-Gregorio (2006) shows that there is a highly convex relationship (again at an aggregate level) between the market value of equity and earnings for German and Italian firms over the period from 1995 to 2005. Moreover, our own unreported empirical work covering 18 broad based UK industrial classifications over the period from 2001 to 2004 shows that the convex relationship obtained at an aggregate level by previous authors also consistently holds at an industry level. Hence, there is now overwhelming international empirical evidence in support of the hypothesis that there is a highly non-linear and generally convex relationship between earnings and the market value of equity.

[7] Ataullah et al. (2006) summarise empirical evidence which is broadly compatible with this value of the normalised risk parameter, θB.

**Figure 1**
**Plot of recursion value of equity, real option value of equity, overall market value of equity and linear approximation to overall value of equity for a branching process with risk parameter θB = 2**




The upward sloping line emanating from the origin at a 45 degree angle is the normalised recursion value of equity, h. The downward sloping curve which asymptotes towards the recursion value axis is the normalised real option value of equity,

$$\frac{1}{2}\int_{-1}^{1}\exp\left(\frac{-4h}{1+z}\right)dz.$$

The sum of the normalised recursion and real option values is the total market value of equity divided by the book value of equity,

$$P(h,1,2) = h + \frac{1}{2}\int_{-1}^{1}\exp\left(\frac{-4h}{1+z}\right)dz,$$

and is represented by the convex curve which asymptotes towards the 45 degree line representing the normalised recursion value of equity. The line emanating from the point 0.3333 on the market value axis is the best linear approximation, $P(h,1,2) \approx 0.3333 + 0.8557h$, to the overall market value of equity.

$h > 2.31$ (although this is not shown on the graph). In other words, for individual firms there will be systematic biases in the linear model which is used to approximate the relationship between the market value of equity and its determining variables.[8]

Here we need to emphasise, however, that the form and nature of the systematic biases which arise from approximating the relationship between equity value and its determining variables in terms of a linear model will very much depend on the magnitude of the normalised risk parameter, θB. The smaller this parameter (and by implication the larger real option values) the more likely it is that a linear model will provide a good approximation to the relationship between the market value of equity and its determining variables. This is illustrated by Figure 2 which shows for a firm whose normalised risk parameter is θB = 0.25, that the best linear approximation to the equity valuation function will be:

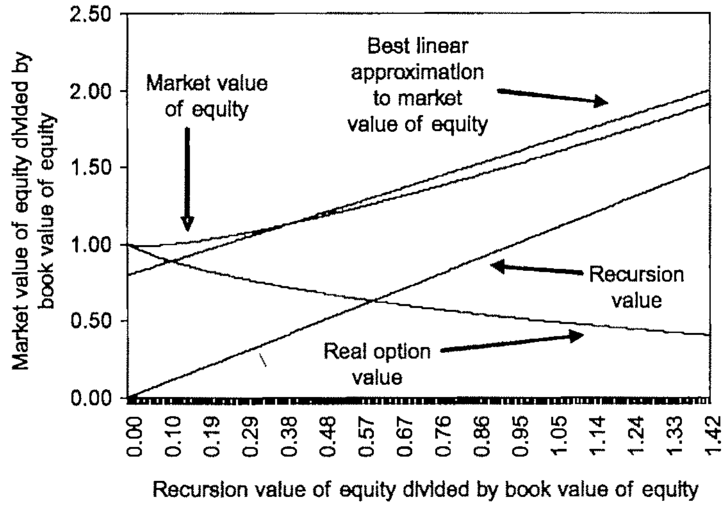
$$P(h,1,0.25) = h + \frac{1}{2}\int_{-1}^{1}\exp\left(\frac{-0.5h}{1+z}\right)dz \tag{10}$$

$$\approx 0.8000 + 0.7976h$$

Note that with a relatively small normalised risk

[8] It is somewhat surprising that linear valuation models continue to dominate analytical and empirical work in this area given that Black and Scholes (1973: 649–652) and Cox and Ross (1976: 163–165) demonstrated over 30 years ago that it is highly unlikely such models can provide an adequate description of the way equity prices evolve in practice.

**Figure 2**
**Plot of recursion value of equity, real option value of equity, overall market value of equity and linear approximation to overall value of equity for a branching process with risk parameter θB = 0.25**




The upward sloping line emanating from the origin at a 45 degree angle is the normalised recursion value of equity, h. The downward sloping curve which asymptotes towards the recursion value axis is the normalised real option value of equity,

$$\frac{1}{2}\int_{-1}^{1}\exp\left(\frac{-0.5h}{1+z}\right)dz.$$

The sum of the normalised recursion and real option values is the total market value of equity divided by the book value of equity,

$$P(h,1,0.25) = h + \frac{1}{2}\int_{-1}^{1}\exp\left(\frac{-0.5h}{1+z}\right)dz,$$

and is represented by the convex curve which asymptotes towards the 45 degree line representing the normalised recursion value of equity. The line emanating from the point 0.8000 on the market value axis is the best linear approximation, $P(h,1,0.25) \approx 0.8000 + 0.7976h$, to the overall market value of equity.

parameter like this it is only at very small and very large ratios of the recursion to book value of equity that linear approximations will provide a poor reflection of the relationship between the market value of equity and its determining variables.

One can further illustrate the importance of the systematic biases demonstrated in these examples by thinking of the valuation equation P(h,1,θB) as a representative firm in a large cross-sectional sample of similarly prepared firms. By 'similarly prepared' is meant that all firms are characterised by a common investment opportunity set and are, therefore, described by a common equity valuation function; namely, P(h,1,θB).[9] It then follows that if equity values include a real option component, cross sectional linear regression models of the relationship between equity prices and recursion values will follow a pattern similar to that obtained for the above examples. That is, one would expect to find firms with low ratios of recursion value to the book value of equity returning negative residuals from a linear regression model. Likewise, firms with intermediate ratios of recursion value to the book value of equity will return positive residuals from the linear regression model. Finally, when the ratio of recursion value to the book value of equity is large, one would expect to see negative residuals again emerging from the linear regression model.[10]

[9] This approach underscores much of the empirical work conducted in the area. Dechow et al. (1999) and Collins, et al. (1999), Morel (2003) and Gregory et al. (2005) provide some recent examples.

## 3. The relative importance of non-linearities in equity valuation

Our previous analysis establishes the inevitability of systematic biases in models that presume a purely linear relationship between the market value of equity and its determining variables. We now investigate whether it is possible to isolate the contribution which the linear and non-linear components of the relationship between equity value and its determining variables make to overall equity value and, in particular, whether it might be possible to characterise equity value in terms of a low order polynomial expansion of its determining variables. We begin by noting that the inner product (Hilbert) space framework employed earlier can be used to express equity value as an infinite series of orthogonal (that is, uncorrelated) polynomial terms of the determining variables and this in turn, allows one to ascertain the relative contribution which each polynomial term makes to the overall variation in equity value. One can illustrate the point being made here by noting that the inner product (Hilbert) space framework implies that the Ashton et al. (2003) equity valuation formula, $P(h,1,\theta B)$, can be expressed in terms of an infinite series expansion of Laguerre polynomials, namely:[11]

$$P(h,1,\theta B) = \sum_{n=0}^{\infty} \alpha_n L_n(h) \tag{12}$$

where $L_0(h) = 1, L_1(h) = 1 - h$ and when $n \geq 2$, $nL_n(h) = (2n - 1 - h)L_{n-1}(h) - (n-1)L_{n-2}(h)$ are the Laguerre polynomials and

$$\alpha_0 = \theta B \log\left(\frac{\theta B}{1+\theta B}\right) + 2,\ \alpha_1 = -1 - \frac{\theta B}{1+\theta B} - \theta B \log\left(\frac{\theta B}{1+\theta B}\right)$$

and when $n \geq 2$,

$$\alpha_n = \frac{\theta B(1+\theta B)^n - (\theta B)^n(n+\theta B)}{n(n-1)(1+\theta B)^n}$$

are the coefficients associated with each of the Laguerre polynomials in the series expansion.

Now suppose one approximates the equity valuation function as a linear sum of the first (m + 1) Laguerre polynomials, or:

$$P(h,1,\theta B) \approx \sum_{n=0}^{m} \alpha_n L_n(h) \tag{13}$$

It can then be shown that:

$$\alpha_0^2 . [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1}$$

gives the proportion of the squared variation in the equity valuation function, $P(h,1,\theta B)$, which is accounted for by the Laguerre polynomial $L_0(h) = 1$ (Apostol, 1967: 566). Likewise,

$$\alpha_1^2 . [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1}$$

gives the proportion of the squared variation in the equity valuation function which is accounted for by the Laguerre polynomial $L_1(h) = 1 - h$. Similarly,

$$\alpha_2^2 . [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1}$$

gives the proportion of the squared variation which is accounted for by the Laguerre polynomial $L_2(h) = \frac{1}{2}(h^2 - 4h + 2)$. Continuing this procedure shows that:

$$R_m^2 = \sum_{n=0}^{m} \alpha_n^2 . [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1} \tag{14}$$

gives the proportion of the squared variation in the equity valuation function which is accounted for by the first (m + 1) Laguerre polynomials. We now show that one can use these results to determine the relative contribution which the linear and non-linear components of the equity valuation function make to overall equity value.[12]

Table 1 summarises the relative contribution which the Laguerre polynomials of order m = 0 to m = 100 make to the overall squared variation of the Ashton et al. (2003) equity valuation function,

---

[10] If one has a large sample of firms then it is not hard to show that under the Ohslon (1995) model a graph of the market value of equity against earnings will, except for a scaling factor, be equivalent to the graph of the market value of equity against its recursion value. Recall here that the Ohlson (1995) model predicts a linear relationship between equity value and earnings in contrast to the highly non-linear and generally convex relationship that exists between these variables in the empirical work summarised by Burgstahler and Dichev (1997), Burgstahler (1998), Ashton et al. (2003) and, Di-Gregorio (2006) amongst others. The convex relationship obtained in this empirical work provides strong indirect evidence that real options have a significant role to play in the determination of equity values.

[11] The proofs for this and subsequent results appear in the Appendix.

[12] The procedure articulated here is equivalent to the spectral decomposition of a variance-covariance matrix using the method of Principal Components (Rao, 1964). The squared coefficients, $\alpha_n^2$, are equivalent to the eigenvalues of the variance-covariance matrix. In the same way as the ratio of a given eigenvalue to the sum of all eigenvalues gives the proportion of the variance accounted for by the eigenvector (principal component) corresponding to the given eigenvalue, the ratio of the squared coefficient to the sum of all squared coefficients gives the proportion of the squared variation in the equity valuation function accounted for by the particular Laguerre polynomial.

**Table 1**
**Proportion of squared variation in equity valuation function associated with Laguerre polynomials**

| | ← $\theta B = 0.25$ → | | | ← $\theta B = 0.50$ → | | | ← $\theta B = 0.75$ → | | | ← $\theta B = 1.00$ → | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| m | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2]^{-1}$ | $R_m^2$ | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2]^{-1}$ | $R_m^2$ | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2]^{-1}$ | $R_m^2$ | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2]^{-1}$ | $R_m^2$ |
| 0 | 1.5976 | 0.798338 | 0.798338 | 1.4507 | 0.768643 | 0.768643 | 1.3645 | 0.739860 | 0.739860 | 1.3069 | 0.714979 | 0.714979 |
| 1 | –0.7976 | 0.198996 | 0.997334 | –0.7840 | 0.224510 | 0.993153 | –0.7931 | 0.249942 | 0.989802 | –0.8069 | 0.272539 | 0.987518 |
| 2 | 0.0800 | 0.002002 | 0.999336 | 0.1111 | 0.004509 | 0.997662 | 0.1224 | 0.005958 | 0.995760 | 0.1250 | 0.006541 | 0.994059 |
| 3 | 0.0373 | 0.000436 | 0.999772 | 0.0617 | 0.001392 | 0.999054 | 0.0758 | 0.002283 | 0.998044 | 0.0833 | 0.002907 | 0.996966 |
| 4 | 0.0203 | 0.000128 | 0.999900 | 0.0370 | 0.000501 | 0.999555 | 0.0491 | 0.000960 | 0.999003 | 0.0573 | 0.001374 | 0.998341 |
| 5 | 0.0124 | 0.000048 | 0.999948 | 0.0239 | 0.000208 | 0.999763 | 0.0333 | 0.000442 | 0.999445 | 0.0406 | 0.000691 | 0.999031 |
| 6 | 0.0083 | 0.000022 | 0.999970 | 0.0164 | 0.000098 | 0.999861 | 0.0236 | 0.000221 | 0.999667 | 0.0297 | 0.000369 | 0.999400 |
| 7 | 0.0060 | 0.000011 | 0.999981 | 0.0118 | 0.000051 | 0.999912 | 0.0174 | 0.000120 | 0.999786 | 0.0223 | 0.000209 | 0.999609 |
| 8 | 0.0045 | 0.000006 | 0.999987 | 0.0089 | 0.000029 | 0.999941 | 0.0132 | 0.000069 | 0.999856 | 0.0172 | 0.000124 | 0.999733 |
| 9 | 0.0035 | 0.000004 | 0.999991 | 0.0069 | 0.000018 | 0.999958 | 0.0104 | 0.000043 | 0.999898 | 0.0136 | 0.000078 | 0.999811 |
| 10 | 0.0028 | 0.000002 | 0.999994 | 0.0056 | 0.000011 | 0.999970 | 0.0083 | 0.000027 | 0.999926 | 0.0110 | 0.000051 | 0.999862 |
| 15 | 0.0012 | 0.000000 | 0.999998 | 0.0024 | 0.000002 | 0.999991 | 0.0036 | 0.000005 | 0.999978 | 0.0048 | 0.000009 | 0.999959 |
| 30 | 0.0003 | 0.000000 | 1.000000 | 0.0006 | 0.000000 | 0.999999 | 0.0009 | 0.000000 | 0.999997 | 0.0011 | 0.000001 | 0.999995 |
| 50 | 0.0001 | 0.000000 | 1.000000 | 0.0002 | 0.000000 | 1.000000 | 0.0003 | 0.000000 | 0.999999 | 0.0004 | 0.000000 | 0.999999 |
| 75 | 0.0000 | 0.000000 | 1.000000 | 0.0001 | 0.000000 | 1.000000 | 0.0001 | 0.000000 | 1.000000 | 0.0002 | 0.000000 | 1.000000 |
| 100 | 0.0000 | 0.000000 | 1.000000 | 0.0001 | 0.000000 | 1.000000 | 0.0001 | 0.000000 | 1.000000 | 0.0001 | 0.000000 | 1.000000 |

**Table 1**
**Proportion of squared variation in equity valuation function associated with Laguerre polynomials** *(continued)*

| | $\theta B = 2$ | | | $\theta B = 4$ | | | $\theta B = 6$ | | | $\theta B = 8$ | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| m | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1}$ | $R_m^2$ | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1}$ | $R_m^2$ | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1}$ | $R_m^2$ | $\alpha_m$ | $\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1}$ | $R_m^2$ |
| 0 | 1.1891 | 0.648684 | 0.648684 | 1.1074 | 0.590307 | 0.590307 | 1.0751 | 0.564577 | 0.564577 | 1.0577 | 0.550209 | 0.550209 |
| 1 | –0.8557 | 0.335968 | 0.984652 | –0.9074 | 0.396343 | 0.986650 | –0.9322 | 0.424505 | 0.989082 | –0.9466 | 0.440686 | 0.990896 |
| 2 | 0.1111 | 0.005664 | 0.990316 | 0.0800 | 0.003081 | 0.989731 | 0.0612 | 0.001831 | 0.990913 | 0.0494 | 0.001199 | 0.992095 |
| 3 | 0.0864 | 0.003426 | 0.993743 | 0.0693 | 0.002314 | 0.992044 | 0.0554 | 0.001499 | 0.992412 | 0.0457 | 0.001028 | 0.993123 |
| 4 | 0.0679 | 0.002115 | 0.995858 | 0.0603 | 0.001748 | 0.993793 | 0.0502 | 0.001230 | 0.993642 | 0.0424 | 0.000883 | 0.994006 |
| 5 | 0.0539 | 0.001333 | 0.997191 | 0.0525 | 0.001329 | 0.995122 | 0.0455 | 0.001013 | 0.994655 | 0.0393 | 0.000759 | 0.994765 |
| 6 | 0.0433 | 0.000858 | 0.998050 | 0.0460 | 0.001016 | 0.996138 | 0.0414 | 0.000836 | 0.995491 | 0.0365 | 0.000654 | 0.995420 |
| 7 | 0.0351 | 0.000565 | 0.998614 | 0.0403 | 0.000782 | 0.996920 | 0.0376 | 0.000692 | 0.996183 | 0.0339 | 0.000565 | 0.995984 |
| 8 | 0.0287 | 0.000379 | 0.998994 | 0.0355 | 0.000606 | 0.997526 | 0.0343 | 0.000575 | 0.996758 | 0.0315 | 0.000488 | 0.996472 |
| 9 | 0.0238 | 0.000260 | 0.999253 | 0.0313 | 0.000472 | 0.997998 | 0.0313 | 0.000479 | 0.997236 | 0.0293 | 0.000423 | 0.996895 |
| 10 | 0.0199 | 0.000182 | 0.999435 | 0.0277 | 0.000370 | 0.998369 | 0.0286 | 0.000400 | 0.997636 | 0.0273 | 0.000367 | 0.997261 |
| 15 | 0.0093 | 0.000040 | 0.999821 | 0.0159 | 0.000121 | 0.999350 | 0.0187 | 0.000170 | 0.998861 | 0.0194 | 0.000185 | 0.998494 |
| 30 | 0.0023 | 0.000002 | 0.999978 | 0.0045 | 0.000010 | 0.999908 | 0.0065 | 0.000021 | 0.999798 | 0.0079 | 0.000031 | 0.999669 |
| 50 | 0.0008 | 0.000000 | 0.999996 | 0.0016 | 0.000001 | 0.999982 | 0.0024 | 0.000003 | 0.999959 | 0.0032 | 0.000005 | 0.999928 |
| 75 | 0.0004 | 0.000000 | 0.999999 | 0.0007 | 0.000000 | 0.999996 | 0.0011 | 0.000001 | 0.999992 | 0.0014 | 0.000001 | 0.999986 |
| 100 | 0.0002 | 0.000000 | 1.000000 | 0.0004 | 0.000000 | 1.000000 | 0.0006 | 0.000000 | 1.000000 | 0.0008 | 0.000000 | 1.000000 |

The above table gives the mth order Laguerre coefficient, $\alpha_m$, and the proportion,

$$\alpha_m^2 \cdot [\sum_{n=0}^{\infty} \alpha_n^2)]^{-1},$$

of the squared variation in the equity valuation function, P(h,1,θB), which is accounted for by the given Laguerre polynomial. It also gives the accumulated proportion, $R^2{}_m$, of the squared variation in the equity valuation function which is accounted for by the first m Laguerre polynomials.

$P(h,1,\theta B)$, for values of the scaled risk parameter that vary from $\theta B = 0.25$ to $\theta B = 8$.[13] Note how the Table shows that a linear approximation ($R_1^2$), based on the coefficients $\alpha_0$ and $\alpha_1$, accounts for over 98% of the squared variation of the Ashton et al. (2003) equity valuation function, irrespective of the scaled risk parameter, $\theta B$, on which the approximation is based. If, for example, one follows the analysis in Section 2 by letting $\theta B = 2$ then Table 1 shows $\alpha_0 = 1.1891$ and $\alpha_1 = -0.8557$ or that consistent with equation (9), the best linear approximation to the equity valuation function is:

$$P(h,1,2) = h + \frac{1}{2}\int_{-1}^{1} \exp\left(\frac{-4h}{1+z}\right)dz$$

$$\approx \alpha_0 + \alpha_1(1-h) = 0.3333 + 0.8557h$$

Moreover, Table 1 also shows that this simple linear approximation accounts for $R_1^2 = 98.4652\%$ of the squared variation in $P(h,1,2)$. This has the important implication that the non-linear terms in the polynomial expansion account for no more than $1 - R_1^2 = 1 - 0.984652 = 1.5348\%$ of the squared variation in $P(h,1,2)$. A similar conclusion applies for other values of the scaled risk parameter, $\theta B$, summarised in Table 1; the non-linear terms in the polynomial expansion make only a minor contribution to the squared variation in the equity valuation function. Given this, one might conclude that a linear approximation of the relationship between equity value and its determining variables will suffice for most practical purposes. Unfortunately, the least squares procedures employed here and also in most of the empirical work of the area, suffer from a ringing artifact known as Gibbs' phenomenon.[14] In the present context Gibbs' phenomenon implies that the Laguerre series expansion will display irregular behaviour in an arbitrarily small interval near the origin (Fay and Kloppers, 2006). This, in turn, means that whilst a linear approximation will provide generally reasonable estimates of the equity valuation function when the recursion value of equity is comparatively large, it will, unfortunately, perform poorly near the origin (that is, when the recursion value of equity is relatively small). It is in this latter circumstance that one will need to include higher order terms from the series expansion if there is to be any prospect of obtaining reasonable approximations to the equity valuation function – something that is borne out by the empirical work summarised in the literature (Burgstahler and Dichev, 1997; Burgstahler, 1998; Ashton et al., 2003 and Di-Gregorio, 2006).

The exact degree to which the Laguerre polynomial series expansion must be carried before one obtains reasonable approximations to the equity valuation function near the origin very much depends on the magnitude of the scaled risk parameter, $\theta B$. Larger values of this parameter will generally require the inclusion of higher order polynomial terms. Here one can again follow the analysis in Section 2 in letting $\theta B = 2$ in which case Figure 3 plots the equity valuation function, $P(h,1,2)$, together with its fifth ($m = 5$), tenth ($m = 10$), fifteenth ($m = 15$) and twentieth ($m = 20$) degree Laguerre polynomial series approximations. Note how the fifth degree Laguerre approximation is particularly poor near the origin and that the tenth and fifteenth degree approximations, whilst an improvement, are still not entirely satisfactory. Indeed, it is only when one employs a twentieth degree polynomial expansion that the approximation to the equity valuation function becomes at all reasonable near the origin. The important point here is that even though linear approximations may appear to be more than reasonable over virtually the entire domain of equity values, nonetheless near the origin (where the recursion value of equity is small) they can be especially poor. This means that mis-specification errors are more likely with samples comprised of firms with comparatively small recursion values – for example, those threatened with administration or which are experiencing other forms of financial distress (Barth et al., 1998). In such instances it is doubtful whether linear models of the relationship between the market value of equity and its determining variables can adequately capture the empirical relationships which exist in the area.[15]

## 4. The biases in simple linear models of equity valuation

Additional insights into the nature of the biases that arise from the simple linear models that pervade this area of accounting research can be obtained by supposing one has a large cross-sectional sample of similarly prepared firms. Let the market value of equity for these $j = 1,2,3,\_\_\_, n$ firms be $P_j = P_j(h,1,\theta B)$ and its associated recursion value be $\eta_j$. It then follows that for each of these firms the relationship between the market value of equity and its recursion value will be:[16]

[13] We have previously noted that Ataullah et al. (2006) summarise empirical evidence which is broadly compatible with these values of the normalised risk parameter, $\theta B$.

[14] The 'least squares' techniques on which much of the empirical work of the area is based can also be formalised in terms of an inner product space with a Euclidean norm and will, as a consequence, also be affected by Gibbs' phenomenon.

[15] Empiricists will often (implicitly) acknowledge the non-linear nature of the valuation relationships that exist in these situations by using dummy variables to allow regression coefficients to vary with so called 'extreme' observations (Barth et al., 1998: 5–9). However, whether this procedure can satisfactorily address the omitted variables problems implicit in their empirical work has yet to be demonstrated.

**Figure 3**
**Plot of polynomial approximations to overall market value of equity for branching process with risk parameter θB = 2**

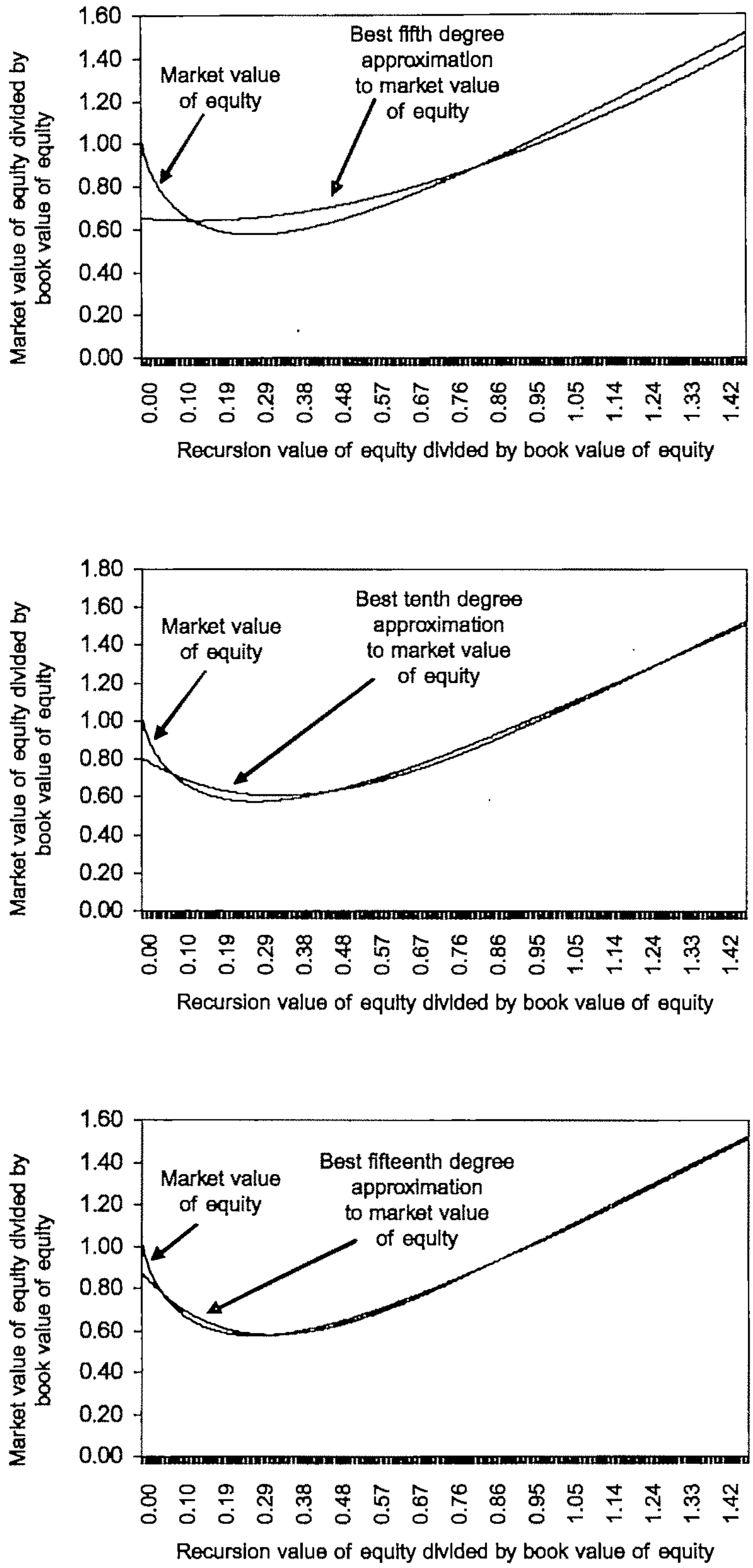
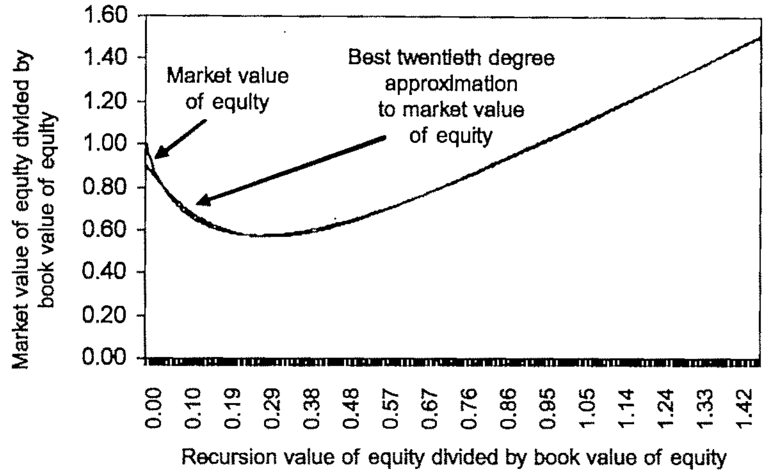


**Figure 3**
**Plot of polynomial approximations to overall market value of equity for branching process with risk parameter θB = 2** *(continued)*




The two curves in the above graphs are firstly, the market value of equity divided by the book value of equity,

$$P(h,1,2) = h + \frac{1}{2}\int_{-1}^{1} \exp\left(\frac{-4h}{1+z}\right)dz.$$

The second curve is the Laguerre polynomial approximation

$$P(h,1,2) \approx \sum_{n=0}^{m} \alpha_n L_n(h)$$

for m = 5 (first graph), m = 10 (second graph), m = 15 (third graph) and m = 20 (fourth graph).

$$P_j = \sum_{k=1}^{m} \beta_k \eta_j^k + e_j \tag{15}$$

where m is the order of the polynomial expansion, $\beta_k$ is the coefficient associated with the $k^{th}$ polynomial term in $\eta_j$ and $e_j$ is the error term associated with the approximation. Now, suppose one fits a linear model, $P_j = \beta\eta_j$, to the given data where $\beta$ is a fixed parameter (Hayn, 1995; Kothari and Zimmerman, 1995). It then follows that the estimate of $\beta$ will be:

$$\hat{\beta} = \frac{\sum_{j=1}^{n} \eta_j . P_j}{\sum_{j=1}^{n} \eta_j^2} = \beta_1 + \frac{\sum_{j=1}^{n} [\beta_2\eta_j^3 + \beta_3\eta_j^4 + \_\_\_\_ + \beta_m\eta_j^{m+1} + \eta_j e_j]}{\sum_{j=1}^{n} \eta_j^2} \tag{16}$$

Hence, if one follows conventional practice in assuming the error term, $e_j$, has a mean of zero then one will obtain the following estimate of $\beta$ on average:

$$E(\hat{\beta}) = \beta_1 + \frac{\sum_{j=1}^{n} [\beta_2\eta_j^3 + \beta_3\eta_j^4 + \_\_\_\_ + \beta_m\eta_j^{m+1}]}{\sum_{j=1}^{n} \eta_j^2} \tag{17}$$

where E(·) is the expectations operator. This result shows that it is unlikely the simple linear models encountered in the literature – of which the empirical work summarised by Dechow et al. (1999), Collins et al. (1999), Morel (2003) and Gregory et al. (2005) are good exemples – can provide dependable information about the relationship between the market value of equity and its recursion value since $\hat{\beta}$ provides a biased estimate of even the first coefficient $\beta_1$, in the series expansion

[16] In this and subsequent analysis pedagogical convenience dictates that we suppress the constant term, $\beta_0$, in the polynomial expansion for $P_j$ without any loss of generality.

for $P_j$.

A more detailed assessment of the biases that are likely to arise from assuming a simple linear relationship between equity value and its determining variables can be made by imposing the conventional assumption that the error term, $e_j$, is an independent and identically distributed normal variate with zero mean and variance, $\sigma^2(e)$. It then follows that the variance of the difference between $\hat{\beta}$ and $\beta_1$ will be:

$$\mathrm{Var}\,(\hat{\beta} - \beta_1) = \mathrm{Var}\left\{ \frac{\sum_{j=1}^{n} \eta_j e_j}{\sum_{j=1}^{n} \eta_j^2} \right\} = \frac{\sigma^2(e)}{\sum_{j=1}^{n} \eta_j^2} \tag{18}$$

One can then compute the standardised variable:

$$z = \frac{(\hat{\beta} - \beta_1) - E(\hat{\beta} - \beta_1)}{\sqrt{\mathrm{Var}(\hat{\beta} - \beta_1)}} \tag{19}$$

which will have a standard normal distribution. Substituting equations (17) and (18) into this expression shows that the standard normal score corresponding to the null hypothesis $\hat{\beta} = \beta_1$ will be:

$$z = -\frac{\sum_{j=1}^{n} [\beta_2 \eta_j^3 + \beta_3 \eta_j^4 + \_\_\_\_ + \beta_m \eta_j^{m+1}]}{\sigma(e) \sqrt{\sum_{j=1}^{n} \eta_j^2}} \tag{20}$$

Previous analysis shows that it is likely equity prices can be represented as a polynomial of order $m \approx 20$ in their determining variables. Given this, under the hypothesis $\hat{\beta} = \beta_1$ it is highly likely that z will be large in absolute terms. This in turn means there is little reason to suppose that the null hypothesis $\hat{\beta} = \beta_1$ would not be rejected at any reasonable level of significance. In other words, there is very little prospect of the regression coefficient, $\hat{\beta}$, in a linear model of the relationship between equity value and its determining variables, providing dependable information about $\beta_1$ – or any other coefficient in the equity pricing relationship for that matter.

## 5. Summary and conclusions

It is now some time since Burgstahler and Dichev (1997: 212) and Penman (2001: 692) observed that the theoretical basis for empirical work on the relationship between equity value and its determining variables is extremely weak. Unfortunately, their call for more refined theoretical modelling in the area has largely been ignored. Empirical work on the relationship between the market value of equity and its determining variables continues to be based on linear models that neglect the real option effects associated with a firm's ability to modify or even abandon its existing operating activities. It is well known, however, that real options induce a convex and potentially, highly nonlinear relationship between equity values and their determining variables (Burgstahler and Dichev, 1997; Ashton et al., 2003). Given this, it is all but inevitable that when real options do impact on equity values, systematic biases will arise in empirical work based on linear valuation models. Our analysis indicates that these biases will be most pronounced for loss making firms. Unfortunately, these firms typically account for around 20% of the samples employed in empirical work (Burgsthaler and Dichev, 1997: 197; Ashton et al., 2003: 428) and so they can have a significant impact on parameter estimation.

Given the now extensive empirical evidence on this convexity issue, it is again timely to renew the call for the development of more refined analytical models of the relationship between equity value and its determining variables. There are two areas in particular where the need for enhanced modelling procedures is urgent. First, relatively little is known about the impact that the real options available to firms have on the book values of assets and liabilities and the accounting policies implemented by firms. The few papers published on this topic (Gietzmann and Ostaszewski, 1999, 2004; Ashton et al., 2003; Ashton et al., 2004) have had relatively little impact on the empirical work conducted in the area – which remains largely wedded to linear valuation models that neglect the impact which real options can have on equity values. Second, even less is known about the appropriate econometric procedures to be used in empirical work in an environment where there is a non-linear relationship between equity values and their determining variables. Suffice it to say that if econometric procedures mistakenly assume the existence of a purely linear relationship between equity values and their determining variables, then it is all but inevitable there will be problems with omitted variables and scale effects in the data. Our analysis shows, however, that these problems can be mitigated by approximating the market value of equity in terms of a polynomial expression of its determining variables – although our evidence is also that the polynomial terms will have to be carried to a fairly high order if this technique is to be satisfactory.

## Appendix
## Polynomial approximation to the equity valuation function

The first two Laguerre polynomials are $L_0(h) = 1$ and $L_1(h) = 1-h$ and these polynomials are orthonormal under the inner product

$$\langle f;g\rangle = \int_0^{\infty} f(h)g(h)e^{-h}dh$$

(Carnahan, et al., 1969: 100). Given this, consider the line of best fit to the equity valuation function in terms of these first two Laguerre polynomials:

$$P(h,1,\theta B) = h + \frac{1}{2}\int_{-1}^{1} \exp\left(\frac{-2B\theta h}{1+z}\right)dz \approx \alpha_0 + \alpha_1(1-h) \tag{A1}$$

where $\alpha_0$ and $\alpha_1$ are known as the Fourier coefficients of the equity valuation equation, $P(h,1,\theta B)$, with respect to $L_0(h)$ and $L_1(h)$, respectively. Now, standard results show (Apostol, 1969: 29–30):

$$\alpha_0 = \langle L_0(h);P(h,1,\theta B)\rangle = \langle 1;P(h,1,\theta B)\rangle = \int_0^{\infty} e^{-h}\left[h + \frac{1}{2}\int_{-1}^{1}\exp\left(\frac{-2\theta Bh}{1+z}\right)dz\right]dh \tag{A2}$$

and:

$$\alpha_1 = \langle L_1(h);P(h,1,\theta B)\rangle = \langle 1-h;P(h,1,\theta B)\rangle = \int_0^{\infty} e^{-h}(1-h)\left[h + \frac{1}{2}\int_{-1}^{1}\exp\left(\frac{-2\theta Bh}{1+z}\right)dz\right]dh \tag{A3}$$

Note, however, that the expression for $\alpha_0$ may be decomposed into two integrals, the first of which is

$$\int_0^{\infty} he^{-h}dh = 1.$$

For the second component, note that all functions under the integral sign are continuous in which case we have:

$$\frac{1}{2}\int_0^{\infty}\int_{-1}^{1} e^{-h}\exp\left(\frac{-2\theta Bh}{1+z}\right)dz\,dh = \frac{1}{2}\int_{-1}^{1}\int_0^{\infty}\exp\left(\frac{-(2\theta Bh + (1+z)h)}{1+z}\right)dh\,dz \tag{A4}$$

where Fubini's Theorem (Apostol, 1969: 363) allows the order of integration to be reversed. One can then evaluate this double integral as follows:

$$\frac{1}{2}\int_{-1}^{1}\int_0^{\infty}\exp\left(\frac{-(2\theta Bh + (1+z)h)}{1+z}\right)dh\,dz = \frac{1}{2}\int_{-1}^{1}\frac{(1+z)}{(2\theta B + (1+z))}dz = \theta B\log\left(\frac{\theta B}{1+\theta B}\right) + 1 \tag{A5}$$

It then follows:

$$\alpha_0 = \langle L_0(h);P(h,1,\theta B)\rangle = \langle 1;P(h,1,\theta B)\rangle = \tag{A6}$$

$$\int_0^{\infty} he^{-h}dh + \frac{1}{2}\int_{-1}^{1}\int_0^{\infty}\exp\left(\frac{-(2\theta Bh + (1+z)h)}{1+z}\right)dh\,dz = \theta B\log\left(\frac{\theta B}{1+\theta B}\right) + 2$$

Similar, though more complicated calculations show that the Fourier coefficient with respect to $L_1(h)$ is:

$$\alpha_1 = \langle L_1(h);P(h,1,\theta B)\rangle = \langle 1 - h;P(h,1,\theta B)\rangle = -1 - \frac{\theta B}{1+\theta B} - \theta B\log(\frac{\theta B}{1+\theta B}) \quad \text{(A7)}$$

Given these results, it follows that the best linear approximation to the equity valuation equation will be:

$$P(h,1,\theta B) \approx \alpha_0 + \alpha_1(1 - h) = [\theta B\log(\frac{\theta B}{1+\theta B}) + 2] + [-1 - \frac{\theta B}{1+\theta B} - \theta B\log(\frac{\theta B}{1+\theta B})](1 - h)$$

or that:

$$P(h,1,\theta B) \approx \frac{1}{1+\theta B} + [1 + \frac{\theta B}{1+\theta B} + \theta B\log(\frac{\theta B}{1+\theta B})].h \quad \text{(A8)}$$

will be the best linear approximation to the valuation function $P(h,1,\theta B)$ over the semi-infinite real line. Non-linear approximations to the equity valuation function can be obtained using the higher order Laguerre polynomials based on the recursion formula (Carnahan, et al., 1969: 100):

$$nL_n(h) = (2n - 1 - h)L_{n-1}(h) - (n-1)L_{n-2}(h) \quad \text{(A9)}$$

where $L_n(h)$ is the Laguerre polynomial of order $n \geq 2$. Moreover, one can use this expression to show that the Fourier coefficient, $\alpha_n$, for the equity valuation function with respect to the nth degree Laguerre polynomial, $L_n(h)$, will be:

$$\alpha_n = \frac{\theta B(1+\theta B)^n - (\theta B)^n(n+\theta B)}{n(n-1)(1+\theta B)^n} \quad \text{(A10)}$$

again provided $n \geq 2$. It then follows that the Ashton et al. (2003) equity valuation formula can be expressed in terms of the following infinite series expansion:

$$P(h,1,\theta B) = \sum_{n=0}^{\infty} \alpha_n L_n(h) \quad \text{(A11)}$$

One can illustrate these latter results by letting $n = 2$ in the recursion formula for the Laguerre polynomials in which case one has that the Laguerre polynomial of order two will be $2L_2(h) = (3 - h)L_1(h) - L_0(h)$ or $L_2(h) = \frac{1}{2}(h^2 - 4h + 2)$. The Fourier coefficient, $\alpha_2$, for the equity valuation function with respect to $L_2(h)$ will then be:

$$\alpha_2 = \langle L_2(h);P(h,1,\theta B)\rangle = \frac{\theta B(1+\theta B)^2 - (\theta B)^2(2+\theta B)}{2.(2-1)(1+\theta B)^2} = \frac{1}{2}\cdot\frac{\theta B}{(1+\theta B)^2} \quad \text{(A12)}$$

It then follows that the best quadratic approximation to the valuation function $P(h,1,B\theta)$ over the semi-infinite real line will be:

$$P(h,1,\theta B) \approx \sum_{n=0}^{2} \alpha_n L_n(h) = \alpha_0 + \alpha_1(1 - h) + \frac{\alpha_2}{2}(h^2 - 4h + 2)$$

or, upon collecting terms:

$$P(h,1,\theta B) \approx \frac{\frac{3}{2}\theta B + 1}{(1+\theta B)^2} + [1 + \frac{(\theta B)^2}{(1+\theta B)^2} + \theta B\log(\frac{\theta B}{1+\theta B})].h + \frac{\frac{1}{4}\theta B}{(1+\theta B)^2}.h^2 \quad \text{(A13)}$$

Moreover, this quadratic approximation to the valuation function $P(h,1,\theta B)$ will be an improvement on the linear approximation summarised earlier. This follows from the fact that the squared error from approximating $P(h,1,\theta B)$ as a linear sum of the first m Laguerre polynomials is given by the squared norm:[17]

[17] A simple proof of this result is given on the Wolfram Mathworld website: http://mathworld.wolfram.com/BesselsInequality.html.

$$\|P(h,1,\theta B) - \sum_{n=0}^{m} \alpha_n L_n(h)\|^2 = \|P(h,1,\theta B)\|^2 - \sum_{n=0}^{m} \alpha_n^2 \tag{A14}$$

where

$$\|P(h,1,\theta B)\|^2 = \langle P(h,1,\theta B);P(h,1,\theta B)\rangle = \int_0^\infty e^{-h}[h + \frac{1}{2}\int_{-1}^{1} \exp(\frac{-2\theta Bh}{1+z})dz]^2 dh$$

is the squared norm of the valuation function and $\alpha_n = \langle L_n(h);P(h,1,\theta B)\rangle$ is the Fourier coefficient of the equity valuation function with respect to the nth order Laguerre polynomial. Since $\alpha_n^2 \geq 0$ for all n it follows from the right-hand side of equation (A14) that the squared error,

$$\|P(h,1,\theta B) - \sum_{n=0}^{m} \alpha_n L_n(h)\|^2,$$

declines as the order of polynomial approximation is increased. Indeed, letting $m \to \infty$ in this expression for the squared error leads to Parseval's relation, namely (Apostol, 1967: 566):

$$\|P(h,1,\theta B)\|^2 = \sum_{n=0}^{\infty} \alpha_n^2 \tag{A15}$$

Now, from equation (A10) when $n \geq 2$ we have:

$$\alpha_n^2 = \frac{\theta^2 B^2}{n^2(n-1)^2} - \frac{2\theta B}{n}.\{\frac{1}{1+(\theta B)^{-1}}\}^n.\frac{(1+\frac{\theta B}{n})}{(n-1)^2} + \{\frac{1}{1+(\theta B)^{-1}}\}^{2n}.\frac{(1+\frac{\theta B}{n})^2}{(n-1)^2} \tag{A16}$$

Consider the last term on the right-hand side of this expression, namely:

$$\{\frac{1}{1+(\theta B)^{-1}}\}^{2n}.\frac{(1+\frac{\theta B}{n})^2}{(n-1)^2} \leq \frac{(1+\theta B)^2}{(n-1)^2} \tag{A17}$$

This in turn means by the Weierstrass M test that the series expansion

$$\sum_{n=2}^{\infty} \{\frac{1}{1+(\theta B)^{-1}}\}^{2n}.\frac{(1+\frac{\theta B}{n})^2}{(n-1)^2}$$

is uniformly and absolutely convergent over the semi-infinite real line (Spiegel, 1974: 228). Similar considerations show that all other terms in the series expansion for

$$\sum_{n=2}^{\infty} \alpha_n^2$$

are uniformly and absolutely convergent over the semi-infinite real line. Given this, one can define the uniformly and absolutely convergent pseudo R-squared statistic:

$$0 \leq R_m^2 = \frac{\sum_{n=0}^{m} \alpha_n^2}{\|P(h,1,\theta B)\|^2} \leq 1 \tag{A18}$$

which gives the proportion of the squared variation in the equity valuation function associated with the best fitting mth order linear combination of Laguerre polynomials. It is then possible to use

$$\alpha_0 = \theta B\log(\frac{\theta B}{1+\theta B}) + 2,\ \alpha_1 = -1 - \frac{\theta B}{1+\theta B} - \theta B\log(\frac{\theta B}{1+\theta B})$$

and for

$$n \geq 2,\ \alpha_n = \frac{\theta B(1+\theta B)^n - (\theta B)^n(n+\theta B)}{n(n-1)(1+\theta B)^n}$$

to evaluate the expression for $R_m^2$ and thereby spectrally decompose the equity valuation function into its various linear and non-linear components.

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# Accounting History

## Call for Papers
## Histories of Accounting Research

Accounting scholars, in general, tend to admire, even if unconsciously, the schools of thought and associated theoretical perspectives that have driven accounting research and structured debates in the academic community and beyond. Often, little is known about the historical development of such schools and the advent and development of key theoretical perspectives which form part of the taken-for-granted underpinnings of frames of reference.

This special issue will comprise articles that provide historical perspectives on schools of thought in accounting. Frequently, examinations of accounting research organise, synthesise and evaluate the published findings of various authors working within a specific paradigm (that is, literature reviews), or undertake a critical analysis of the assumptions and/or methods employed within particular paradigms, such as capital markets research. Accounting History has tended to ignore the research enterprise and focus on specific episodes, such as histories of standard-setting, histories of accounting and auditing techniques and practices, or histories of the profession, firms or prominent individuals. Furthermore, few historical studies in accounting examine the development of accounting research across space. To redress this omission, relevant manuscripts bringing new historical insights about accounting research are cordially invited for review.

Topics may include, but are not limited to, the following areas:

- Examination of changing conceptions of the role of the academic as research paradigms shift, and/or in response to changing constellations of users/supporters of accounting research.
- Identification and evaluation of 'significant' past literature reviews.
- The translation or mutation of research issues as they move across research paradigms. That is, how have 'old' research questions been transformed or resurrected through the application of different research methods?
- Tracing the trajectory of research questions as new research methods are introduced. What questions become possible, and which are abandoned, as research methods shift within a particular research stream such as auditing?
- Identification of abandoned branches of the 'family' tree and identification of any research streams which have become 'locked in'.
- Exploration of the crucial turning points that launched a literature or changed its questions.
- Studies of the relationship(s) between the evolution of accounting research and broader social discourses and the absorption of accounting discourses within other disciplines.
- Studies which explore why certain countries appear to become home base for particular types of research, such as investigations as to why capital markets research, for example, is more prominent in the US and accounting research within the critical paradigm is more widespread in the UK.
- Longitudinal studies of the relationship(s) between accounting, business and economic history.
- Explorations of the role played by accounting history research in broadening our understandings of contemporary accounting as a social and institutional practice.

Potential contributors are encouraged to interpret this theme using diverse theoretical and methodological perspectives and are strongly encouraged to contact the guest editors in advance to discuss their proposed topics. Submissions must be written in English and forwarded electronically, to the guest editors, by 31 October 2009. This special issue is scheduled to be published in late 2010/early 2011.

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Does graph disclosure bias reduce the cost of equity capital?

Value-relevance of presenting changes in fair value of investment properties in the income statement: evidence from Hong Kong

Earnings quality in ex-post failed firms

Selection bias and the Big Four premium: new evidence using Heckman and matching models

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# Forthcoming Articles

The next issue of the journal will be the Special Issue in collaboration with the Institute of Chartered Accountants in England and Wales, containing papers and discussions arising from the PD Leake Lecture Series (2007, 2008) and the International Accounting Policy Forum in December 2008. The major papers will be as follows:

**PD Leake Lecture Series**
Chris Cowton, 2007, 'Accounting and the ethics challenge: re-membering the professional body'
Geoff Meeks and Peter Swann, 2008, 'Accounting standards and the economics of standards'

**International Accounting Policy Forum**
Nicholas Barr, International trends in pension provision
Christopher Napier, The logic of pension accounting
Vicky Kiosse and Ken Peasnell, Have changes in pension accounting changed pension provision? A review of the evidence
Martin Glaum, Pension accounting and research: a review

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All papers are subject to peer review on a double blind basis, either by members of the Editorial Board, or by invited reviewers of international standing. Reviewers are asked to comment in particular on the contribution, motivation and rigour of the analysis presented in the paper. The editor carries out an initial check that papers submitted comply with the guide to authors and advises authors where a paper has not met the essential criteria. Continuous monitoring of the review process aims at providing timely but informative feedback to authors. Subject to the recommendation of reviewers, research notes and commentaries may be published.

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# Does graph disclosure bias reduce the cost of equity capital?

**Flora Muiño and Marco Trombetta***

**Abstract**—Firms widely use graphs in their financial reports. In this respect, prior research demonstrates that companies use graphs to provide a favourable outlook of performance, suggesting that they try to manage the impression created in users' perceptions. This study tests whether by means of distorted graphs managers are able to influence users' decisions in the capital market. By focusing on the effects of distorted graphs on the cost of equity capital, we provide preliminary evidence on one of the possible economic consequences of graph usage. The results of this investigation suggest that graph disclosure bias has a significant, but temporary, effect on the cost of equity. Moreover, our results highlight the important role played by the overall level of disclosure as a conditioning factor in the relationship between graphs and the cost of equity. Consequently, the results of the current study enhance our understanding of the complex interactions that take place in the stock market between information, information intermediaries and investors.

**Keywords:** cost of equity, disclosure, graph distortion

## 1. Introduction

Prior research on graph usage in annual reports has documented two important facts. First, graph usage increases when company performance improves (e.g. Steinbart, 1989 and Beattie and Jones, 1992). Second, an important proportion of graphs are distorted to portray a more favourable view of the company than reflected in the financial statements (e.g. Beattie and Jones, 1999 and Mather et al., 1996). Thus, the existing literature supports the notion that managers use distorted graphs to manage the impression created in the perceptions of users of annual reports[1] (e.g. Beattie and Jones, 2000). However, Beattie and Jones (2008: 22), after reviewing the existing literature on graph usage in annual reports, conclude:

'A fundamental issue to be addressed by future research is whether an impact on a user's perceptions of the organisation carries through to an impact on their decision. For example, are investment decisions affected by graphical presentation choices? We need, therefore, to carry out research into the effect, if any, of financial graphs on analysts' earnings forecasts, stock prices and the relationship decisions of other key stakeholder groups such as employees, customers and suppliers'

From a theoretical point of view, the potential effect of graph distortion on the functioning of the stock market is a controversial issue. On the one hand, the literature on 'herding' and 'limited attention' in financial markets shows that economic agents may neglect a considerable part of their private information because of reputation and coordination effects or information processing limitations.[2] Consequently, graphs can have an impact on users' judgments and decisions, even when the information represented in graphs can be gathered also in other parts of the annual report. On the other hand, the efficient markets hypothesis (EMH) tells us that prices should include all the publicly available information. Hence, graph distortion should not affect prices because they should be based on all the information available on

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This paper was accepted for publication in November 2008.

[1] Impression management techniques in annual reports are not limited to the use and distortion of graphs. Extant research provides evidence consistent with the use of narratives as an impression management technique (e.g. Clatworthy and Jones, 2003; Aerts, 2005).

[2] A more detailed review of this literature is provided in the next section of the paper.

the annual report and not only on the image portrayed by graphs. Our aim is to provide empirical evidence that improves our understanding of this controversial issue and take a step forward in the direction indicated by Beattie and Jones (2008).

As an indicator of the capital market effects of graph distortion we use the cost of equity capital, measured both on an ex-ante (i.e. expected) and an ex-post (i.e. realised) basis.[3] The ex-ante measure is based on analysts' forecasts of earnings issued at a certain point in time, soon after the publication of the annual report. This measure depends directly on the expectations of those analysts following the company at the time forecasts are issued. Based on the evidence provided by literature on herding and limited attention, we hypothesise that these expectations can be biased because of distorted graphs included in annual reports. However, on the grounds of the EMH, we do not expect graph distortion to affect our ex-post measure of the cost of equity, realised returns. Since realised returns aggregate the decisions taken by investors over a relatively large period of time (one year), we expect that the potential bias in analysts' perceptions is corrected in the market by means of the aggregation process and the passing of time that allows new information to be impounded in stock prices.[4]

We base our analysis on a sample of Spanish companies quoted in the Madrid Stock Exchange (MSE) between 1996 and 2002. As compared to the US or the UK, where most prior studies on graphs have been developed, Spain has still a relatively underdeveloped capital market, where levels of corporate transparency are generally low.[5] Moreover, prior research documents the existence of a higher level of herding in Spain than in countries with more developed capital markets.[6] If graph distortion is able to bias the perceptions of market participants, that is more likely to be observed in those markets, such as the MSE, with low corporate transparency and high levels of herding. When comparing US and European stock markets (excluding the UK), Bagella et al. (2007) find that the absolute bias in earnings forecasts is significantly higher in Europe than in the US. This is why we believe that the MSE is well suited to investigate whether graph distortion affects perceptions and decisions of market participants.

After controlling for other possible determinants of the cost of equity, we detect a significant negative effect of favourable graph distortion on our ex-ante measure of the cost of equity. This effect is moderated by the overall level of disclosure so that, at high levels of transparency, the relationship between graph distortion and the ex-ante cost of equity becomes positive. However, when we turn to the ex-post analysis, we do not find any significant effect of graph distortion on realised returns. We interpret this result as evidence that the effect on the ex-ante cost of equity is only temporary and confined to expectations. It is really an effect on the bias of the analysts' forecasts that are used to calculate the ex-ante measure of the cost of equity. However, with the passing of time the aggregation process performed in the market corrects this bias.

Our contribution to the literature is twofold. First, to the best of our knowledge, we are the first to provide empirical evidence on the capital market effects of graph distortion for quoted firms. Graph distortion is an ideal item from which to study the economic effects of impression management techniques. As stated by Merkl-Davies and Brennan (2007), the study of these effects is complicated by the difficulty of separating impression management (opportunistic distortion of information) from incremental information (provision of useful additional information). However, graph distortion, as an impression management technique, is not affected by this problem. Given that the information portrayed in graphs is always available in another format in the same annual report, we can assume that distorted graphs are pure impression management tools and do not provide any incremental information.

Second, we provide additional evidence that supports the necessity of distinguishing between estimations by individuals and aggregated market behaviour, when studying the capital market effects of impression management techniques. Short-term effects on individuals' predictions can be corrected with the passing of time by the aggregation role played by market activity, so that they do not have an impact on long-window stock returns.

The structure of the paper is as follows. In the next section we review the existing literature and develop our research hypotheses. In Section 3 we describe our sample and the variables that we use in our empirical analysis. Section 4 contains the

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[3] A review of the issues involved in the empirical measurement of the cost of equity capital can be found, for example, in Botosan (2006).

[4] Temporary differences between ex-ante and ex-post measures of the cost of equity are depicted in Figure 1 and discussed in detail in section 3.2.1.

[5] According to the World Federation of Exchanges, in 2002, the latest year covered in our analysis, total value of share trading in the MSE was $653,221m as compared to $4,001,340m for the London Stock Exchange or $10,310,055m for the New York Stock Exchange. Additionally, La Porta et al. (2006) provide evidence showing that the MSE has a lower index of disclosure requirements (0.50, as compared to 0.83 and 1.00 for the UK and the US, respectively) and public enforcement (0.33 as compared to 0.68 and 0.90 for the UK and the US, respectively).

[6] Ferruz Agudo et al. (2008) report a level of herding of 13.26% for Spain, a figure similar to that observed by Lobao and Serra (2002) for Portugal, but much higher than the 2% found by Lakonishok et al. (1992) for the US or the 3.3% observed by Wylie (2005) in the UK.

main results of our analysis. Section 5 summarises the analysis, provides some conclusions and draws implications.

## 2. Background and hypotheses development

The importance of information for the functioning of the stock market is manifest. The whole notion of market efficiency in its different forms is based on how agents in the market react to information and how information is eventually reflected into prices. Graphs included in annual reports do not provide additional information, but they are a vehicle of information dissemination that makes data more visible and, when properly constructed, facilitates its processing (Beattie and Jones, 1992). Extant research, however, shows that a significant proportion of graphs included by companies in their annual reports is materially distorted, generally to portray corporate performance more favourably[7] (e.g. Mather et al., 1996; Beattie and Jones, 1999). Researchers conclude that managers use distorted graphs to manage users' perceptions (e.g. Beattie and Jones, 2000).

Experimental research provides evidence consistent with this conclusion. Taylor and Anderson (1986), Beattie and Jones (2002) and Arunachalam et al. (2002) all provide evidence that users' perceptions are far more favourable when they are based on graphs that are favourably distorted. Arunachalam et al. (2002) explain this phenomenon on the basis of two findings of psychological research. First, Payne et al. (1993) show that people minimise their cognitive effort in order to achieve a certain level of accuracy. Second, Ricketts (1990) finds that people have difficulty at detecting presentation errors. Distorted graphs usually do not reverse or fundamentally change the reality as expressed by the data.

Participants' perceptions are affected by distorted graphs even when the accurate numeric values are displayed as variable labels in graphs (e.g. Arunachalam et al., 2002). This suggests that participants focus their attention on the image portrayed in graphs and ignore the numeric values. The fact that participants in these experimental studies neglect relevant information could be attributed to their low level of experience (they were students). However, extant research provides evidence showing that even sophisticated users (analysts) do not always make use of all available information.

A substantial body of research in theoretical finance demonstrates that, under certain circumstances, analysts may ignore a significant proportion of their private information. Trueman (1994) demonstrates that analysts tend to release forecasts similar to those reported by other analysts, even when their private information does not justify such forecasts. Similarly, Morris and Shin (2002) analyse the role of public information in contests where an agent needs to co-ordinate with other agents in order to maximise his/her payoff.[8] They show that, in these settings, it can be socially optimal to adopt coarser information systems instead of finer information systems. These theoretical predictions have been confirmed empirically. Stickel (1992), Graham (1999), Hong et al. (2000), and Welch (2000) provide evidence of analysts' herding behaviour and Anctil et al. (2004) use an experimental setting to document the negative social effects of a lack of coordination.

Literature on limited attention provides empirical evidence showing that investors ignore valuable information when making investment decisions. For instance, Doyle et al. (2003) find that the stock market does not fully appreciate the predictive power of expenses excluded from proforma earnings. Similarly, Hirshleifer et al. (2004) demonstrate that investors do not optimally use the information conveyed by net operating assets when assessing the sustainability of corporate performance. Furthermore, experimental studies such as those by Hopkins (1996), Hirst and Hopkins (1998) and Hirst et al. (2004) show that analysts' judgments and valuations are affected by the income-measurement method (recognition versus disclosure in the footnotes) or the classification of items in the financial statements. These studies provide further evidence showing that even experts ignore relevant information when making estimations and valuations; otherwise their valuations would not be affected by classification of information in financial statements. Overall, prior research indicates that users tend to focus their attention on the most salient and easily processed information, neglecting relevant data. This is attributed to limited attention and cognitive processing power (Hirshleifer and Teoh, 2003).

Graphs represent a prominent piece of information in the annual report and the information they convey can be processed by users fairly easily. Then, on the basis of both the theoretical arguments and the empirical and experimental evidence provided by extant literature, we hypothesise that market expectations and analysts' estimations can be affected by favourable graph distortion, despite the fact that accurate values of the variables dis-

[7] Distortion refers to violations of what Tufte (1983) states as an essential principle in graph construction: physical measures on the surface of the graph should be directly proportional to the numerical quantities represented. The use of non-zero axis, broken axis, or non-arithmetic scales leads to graphs where equal distances along the axis do not represent equal amounts, that is, physical measures are not proportional to the underlying numerical values.

[8] The stock market is used by Morris and Shin (2002) as an example of such a contest.

played in graphs are reflected in the financial statements. The ex-ante (expected) measures of the cost of equity capital depend directly on these expectations and estimations in the sense that more optimistic forecasts about the future of the company are associated with a lower level of the ex-ante cost of capital. Hence, we state the following hypothesis to be tested in our study:

**H1:** Favourable graph distortion is negatively related to the ex-ante cost of equity capital.

Disclosure and time are two additional factors that can affect the relationship between graph distortion and the cost of equity capital. Next, we discuss the moderating effect that can be exercised by disclosure and afterwards we analyse the role played by the passing of time in correcting biases in ex-ante expectations.

The above-mentioned prior experimental research indicates that users have difficulties in detecting graph distortion. Corporate disclosure can ameliorate this situation by providing users with additional data that may turn out to be important clues in identifying distorted graphs.[9] Take, for example, the case of a firm presenting a slight increase in net sales, which is magnified in a distorted graph so that it appears as a strong rising trend. Distortion is difficult to detect because the direction of the change depicted graphically is the same as that shown by data in financial statements. However, if this firm provides operating data showing a decrease in production, readers are more likely to realise the graph is distorted. The role played by disclosure in correcting misperceptions is documented in the literature. Schrand and Walther (2000) observe that the bias introduced by strategic choices of prior-period benchmarks in earnings announcements is eliminated when the financial statements are released. Along the same line, in a controlled experiment, Krische (2005) finds that clear and quantitative information about prior-period transitory gains or losses allows participants to adjust the comparative prior period-earnings stated as a benchmark in earnings announcements. Based on prior evidence we expect that the likelihood of distortion being detected is increasing in the level of disclosure.

Detection of graph distortion can have an impact on users' decisions because of the impairment in corporate disclosure credibility. As stated by Schmid (1992), graph distortion threatens the credibility of the entire report containing such a graphic. Therefore, when distortion is detected, users may perceive a higher risk associated with their decisions.[10] Theoretical studies by Easley and O'Hara (2004) and Leuz and Verrecchia (2005) predict the existence of a negative relationship between the quality of information and the risk premium required by investors. Easley and O'Hara (2004) show that information precision reduces the information-based systematic risk of shares to uninformed investors, thereby reducing the cost of capital. Leuz and Verrecchia (2005) take a different approach and show that information quality increases expected cash-flows and, as a consequence, reduces the firms' cost of capital. Francis et al. (2004) and Francis et al. (2005) provide empirical evidence supporting these predictions. They find that accrual quality and a number of earnings attributes are significantly related to the cost of capital. By analogy we expect that, by reducing the credibility of the annual report, the inclusion of distorted graphs, if detected, will increase the information risk and, as a consequence, the risk premium demanded by investors. On the contrary, as long as distortion is undetected, companies could benefit from a lower cost of equity. Since the overall level of disclosure can be essential in unveiling graph distortion, we test the following hypothesis:

**H2:** The effect of favourable graph distortion on the ex-ante cost of equity capital is different at low levels of disclosure than at high levels of disclosure.

So far we have focused our attention on users' perceptions and estimations about the future of the company. These perceptions and estimations are at the basis of the so called ex-ante measures of the cost of equity. After estimations have been formed and revealed, investors will take their buying and selling decisions in the stock market and market prices will be formed continuously. The EMH predicts that market prices reflect all available information in an efficient manner. Hence, it could be argued that graph usage and graph distortion should not affect the aggregated ex-post behaviour of the stock market. In other words, as time passes, the aggregation role played by market activity should correct any possible bias contained in the ex-ante perceptions and estimations.

In a paper closely related to our analysis, Easton and Sommers (2007) provide evidence showing that ex-ante measures of the cost of equity are upwards biased because of the bias contained in analysts' forecasts. However, they also show that the

[9] The relationship between the overall level of disclosure and the cost of equity has already been explored in various papers, both from a theoretical and an empirical point of view. For a general introduction to this literature see Botosan (2006). A more analytical perspective can be found in Verrecchia (2001) and Dye (2001).

[10] Prior research referring to managers' explanations for poor performance supports this line of reasoning. Barton and Mercer (2005) find that implausible explanations harm management reputation leading to an increase in the firm's information risk. Although these explanations are given to promote a more favourable view of corporate performance, they end up having the contrary effect.

stock market undoes the bias of analysts' forecasts. Hence, these authors suggest that variation in the proxies used to measure the ex-ante cost of equity capital can be due to variation in analysts' bias rather than to variation in the true cost of equity capital. Starting from this result, we conjecture that the bias in analysts' estimations of earnings per share (the base for the ex-ante measure of the cost of equity) can be influenced by graph distortion, while the ex-post realised cost of equity is not. In other words, stock prices are not biased by distorted graphs. This argument is the basis of our third hypothesis.

**H3:** Favourable graph distortion does not affect stock market returns, i.e. the ex-post cost of equity capital.

Taken together, our three research hypotheses aim at covering some important aspects of the potential role played by graphs as communication tools in the stock market. In the next section we implement our research design by describing the sample and the variables used for our empirical tests.

## 3. Research design and variable definitions

### *3.1. The sample*

For our sample we collect information on companies listed on the continuous (electronic) market in the MSE for the period 1996–2002. Information on the graphs and their characteristics is gathered from corporate annual reports. Data on the information disclosed by the company (disclosure index) is obtained from the rankings produced by a pool of analysts and published annually by the business magazine Actualidad Económica.[11] Analysts' forecasts and market data used to calculate the cost of equity capital measures are obtained from the JCF database.[12] Finally, financial data are collected from the OSIRIS database.[13] Companies were removed from our study when the necessary information was not available at least for two consecutive years. Our final sample comprises 259 firm-year observations from 67 companies during 1996–2002.

### *3.2. Definition of variables*

#### *3.2.1. Cost of capital measures*

*Ex-ante cost of capital ($R_{PEG\_PREM}$)*
An ex-ante measure of cost of equity capital is a measure based on some valuation model and it infers the discount rate that it should have been used to calculate the observed price of the stock if that particular valuation model had been used. The question of which is the best proxy for the ex-ante cost of equity capital of a company has received a great deal of attention in the recent literature. To produce a comprehensive review of this literature is beyond the scope of this study. However, one of the common results of these studies is that all the measures proposed in the literature are highly correlated among each other. Consequently, the choice of a particular proxy instead of another should not have a major effect on the overall results of the study. Botosan and Plumlee (2005) assess the relative merits of five alternative proxies in terms of their association with credible risk proxies. They find that the measure proposed by Easton (2004) is one of the two measures that clearly dominate the other three. They observe that it correlates with a number of risk indicators in the expected direction. This result, combined with the relative simplicity of this measure, made us decide to use it as our primary proxy for the ex-ante cost of equity capital.[14] Following Easton (2004) we calculate the ex-ante cost of equity capital for year $t$ as[15]

$$r_{peg} = \sqrt{\frac{eps_2 - eps_1}{P_0}}$$

where:

$P_0$ = price of the stock of the company at 30 June of year $t$+1

$eps_1$ = one year ahead consensus forecast of earnings per share at 30 June of year $t$+1

$eps_2$ = two years ahead consensus forecast of earnings per share at 30 June of year $t$+1

Following Botosan and Plumlee (2005) we then calculate the equity risk premium ($R_{PEG\_PREM}$) by subtracting from the cost of equity capital the risk free rate, proxied by the interest rate on five-year Spanish Treasury bills.[16] To avoid an unduly effect of outliers, $R_{PEG\_PREM}$ is then winsorised at the 1 and 99 percentiles of its distribution (i.e. values in the top and bottom 1% of the distribution are set equal to next value counting inwards from the extremes).

[11] The process of elaboration of this index is similar to that followed in constructing the AIMR index and it is discussed in detail in Section 3.2.3.

[12] JCF provides a global database of consensus earnings estimates and other financial projections to the professional investment community.

[13] The OSIRIS database, compiled by Bureau Van Dijk, provides financials for the world's publicly quoted companies from over 130 countries. It has been used in prior studies (e.g. García Lara et al., 2007; Surroca and Tribó, 2008). After comparing a number of databases, including OSIRIS, García Lara et al. (2006) conclude that when using the same set of companies results are not affected by the choice of the database.

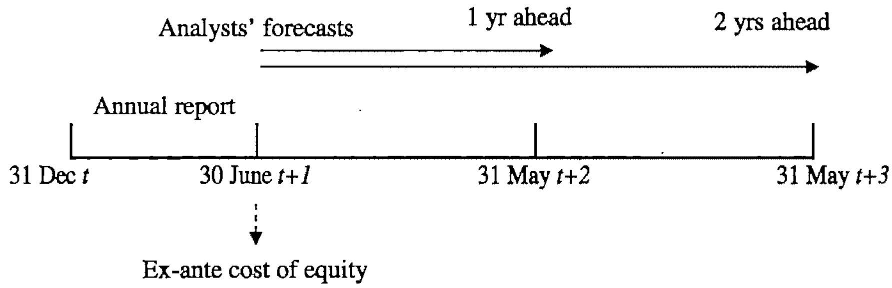
[14] We check the robustness of our results to the use of other proxies for the ex-ante cost of equity. Results of these sensitivity analyses are discussed in Section 4.2.1.

[15] For the exact derivation of the formula the reader can refer to Easton (2004).

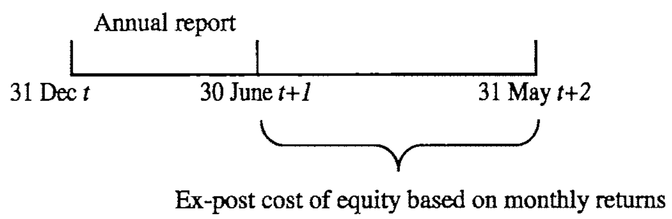
[16] We subtract the risk-free rate to obtain an indicator of the premium for risk required by investors to fund the company.

**Figure 1**
**Ex-ante and ex-post cost of equity**

**Panel A: Ex-ante cost of equity**




**Panel B: Ex-post cost of equity**




*Ex-post cost of capital (stock returns)*
Besides the ex-ante measure of the cost of equity, we use the one-factor asset-pricing model to assess the actual effect of graphs on the aggregated market.[17] Similar to Francis et al. (2005), when analysing the effect of accruals quality in stock returns, we add a variable capturing the level of graph distortion to the traditional capital asset pricing model (CAPM). To obtain this graph distortion factor we use a similar procedure to that employed by Fama and French (1993) in constructing the size and book-to-market factor-mimicking portfolios. We start by constructing two mimicking portfolios for graph distortion. To obtain these portfolios we divide stocks into two groups: distorting and non-distorting companies. Distorting companies are those which materially distort the graphs included in their annual reports to portray a more favourable view of the company.[18] Non-distorting companies are those that present fairly constructed graphs or distorted graphs presenting a more unfavourable view of the company. We calculate monthly excess returns for companies in each group from June of year t+1 to May of year t+2.[19] The graph distortion factor mimicking portfolio equals mean monthly excess return for distorting companies portfolio less mean monthly excess return for non-distorting companies portfolio. For our sample period (1996–2002) we obtain a series of 84 monthly returns for the graph distortion factor. Then, we estimate the CAPM including the graph distortion factor for each of the 67 companies included in our sample. In these regressions, the coefficient of the graph distortion factor indicates whether this factor adds to market risk premium in explaining returns, our proxy for the ex-post cost of equity.

The ex-ante and ex-post measures of the cost of equity differ not only in the nature of the information on which they are based (analysts' forecasts as opposed to realised returns), but also in the period of time when they are formed. The relationship between the ex-ante and the ex-post measure of the cost of equity capital in terms of time is represented in Figure 1.

As we can see, the ex-ante measure is based on analysts' forecasts of earnings per share one and two years ahead at 30 June of year t+1, whereas the ex-post measure is based on realised returns (i.e. we estimate firm-specific asset-pricing models running from June 1997 to May 2004). For the period June of year t+1 to May of year t+2, returns

[17] Similar results obtain when using the three-factor model proposed by Fama and French (1993).

[18] Material distortion refers to a mean favourable graph distortion index above 2.5%. The selection of this cut-off point is explained in detail in Section 3.2.2, when describing the measure of graph distortion. The direction of the results remains unchanged if distorting and non-distorting portfolios include companies in the top and bottom 40% percentiles of the distribution of the variable used to measure graph distortion, respectively.

[19] Following prior literature on disclosure in Spain (e.g. Espinosa and Trombetta, 2007), calculations start in June of year t+1, given that financial reporting regulation in Spain requires companies to release their annual report by that date at the latest.

**Table 1**
**Measures of graph distortion**

**Panel A: Distortion measure for individual graphs**

| *Trend in data* | *Nature of distortion* | *RGD* |
|---|---|---|
| Increasing | Exaggeration | >0 |
| Decreasing | Understatement | >0 |
| Increasing | Understatement | <0 |
| Decreasing | Exaggeration | <0 |

**Panel B: Distortion measure across all graphs in the annual report**

| *Graphs in the annual report* | *RGDFAV* | *RGDUNF* |
|---|---|---|
| All graphs are properly constructed | 0 | 0 |
| There are favourably distorted graphs<br>There are unfavourably distorted graphs | >0 | >0 |
| There are favourably distorted graphs<br>There is not any unfavourably distorted graph | >0 | 0 |
| There is not any favourably distorted graph<br>There are unfavourably distorted graphs | 0 | >0 |

to the graph distortion factor incorporated in the asset pricing models are calculated using the level of graph distortion in the annual report of year t. Then, when using the ex-ante measure of the cost of equity, the analysis of the effect of graph distortion is based on expectations observed at a certain point time (June of year t+1). However, when employing the ex-post measure, the analysis is based on monthly realised returns covering one year (from June of year t+1 to May of year t+2). Accordingly, the latter measure allows for two types of corrections: those arising from the aggregation of all investors' decisions and those resulting from the passing of time which permits new information to be impounded in stock prices.

### *3.2.2. Graph distortion measures*

*Relative graph discrepancy (RGD) index*

Graph distortion is measured in our study by using the relative graph discrepancy (RGD) index (Mather et al.,2005),[20] which is defined as

$$RGD = \frac{g_2 - g_3}{g_3}$$

where $g_2$ represents the actual height of the last column and $g_3$ is the proportionately correct height of the last column, based on the formula:

$$g_3 = \frac{g_1}{d_1} d_2$$

$d_1$ = value of first data point (corresponding to the first column)

$d_2$ = value of last data point (corresponding to the last column)

$g_1$ = actual height of first column

$g_2$ = actual height of last column.

In the absence of distortion, the index takes the value of zero (0), that is, the change portrayed in the graph is the same as that observed in the data. The RGD takes a positive value both when an increasing trend is exaggerated and when a decreasing trend is understated. Negative values result from understatement of increasing trends and exaggeration of decreasing trends (Table 1, Panel A).

*Measures of favourable and unfavourable graph distortion*

The RGD index gives us an indication of the level of distortion of a particular graph, either if distortion is favourable or unfavourable to the firm. To test our hypotheses we need to isolate those distortions that are favourable to the company and design an indicator of favourable (unfavourable) graph distortion across all graphs in the annual report.

[20] This measure was developed by Mather et al. (2005) to overcome some of the limitations of the graph discrepancy index (GDI), the measure of graph distortion used in previous studies. The GDI is defined as $(\frac{a}{b}-1)\times 100$, where $a$ is the percentage of change in centimetres depicted in the graph and $b$ is the percentage of change in the data. We repeated all our analysis using the GDI, instead of the RGD, and results do not vary.

Favourably distorted graphs are those graphs manipulated to present a more favourable view of the company. Examples of favourable distortion are the magnification of a positive trend in sales growth or the understatement of a decreasing trend in the same variable. Conversely, understatement of an increasing trend in sales and exaggeration of a decreasing trend in this variable are examples of unfavourable distortion.

We measure the level of favourable distortion across all graphs included by a firm in the annual report as follows:

$$RGDFAV = \frac{\sum_{j=1}^{n} |rgd\,fav_j|}{n}$$

where:

$|rgd\,fav_j|$ = Absolute value of the RGD index for graph $j$ in the annual report.[21] $rgd\,fav_j$ is set to zero when graph $j$ is distorted to portray a more unfavourable view of the company.

$n$ = total number of graphs in the annual report.

The RGDFAV provides us with an indication of the mean level of favourable graph distortion in the annual report. This measure is increasing in the number of favourably distorted graphs and the corresponding RGD indices. Zero (0) value for the RGDFAV indicates that the annual report does not contain any favourably distorted graph. Panel B in Table 1 describes the values that correspond to the RGDFAV measure depending on whether the annual report includes properly constructed, favourably, and unfavourably distorted graphs.

Similarly, we measure unfavourable graph distortion across all graphs in the annual report as follows:

$$RGDUNF = \frac{\sum_{j=1}^{n} |rgd\,unf_j|}{n}$$

where:

$|rgd\,unf_j|$ = Absolute value of the RGD index for graph $j$ in the annual report. $rgd\,unf_j$ is set to zero when graph $j$ is distorted to portray a more favourable view of the company.

The RGDUNF is increasing in the number and the RGD of distorted graphs presenting a more unfavourable image of the company and takes the value of zero (0) when the annual report does not include any unfavourably distorted graph (Table 1, Panel B).

We also calculate mean favourable distortion index for financial (RGDFFAV) and non-financial graphs (RGDRFAV). These indices are defined in the same way as the RGDFAV but taking into consideration exclusively financial graphs (i.e. graphs depicting financial variables) for the RGDFFAV and non-financial graphs for the RGDRFAV.

In estimating the regression models we use the fractional ranks of our graph distortion measures (RGDFAV, RGDUNF, RGDFFAV and RGDRFAV). Fractional ranks for each of these measures are computed by dividing, within each year, the rank of a firm's distortion measure by the number of firms in the sample in this year. The rank is increasing in the level of distortion.

Finally, in order to distinguish between materially and non-materially distorted graphs we have to choose a cut-off point for the RGD measure. Mather et al. (2005) conclude that an RGD of 2.5% would be similar to a GDI of 5%, the cut-off point suggested by Tufte (1983) and used in previous studies. This is why we decided to use the 2.5% cut-off point as suggested by Mather et al. (2005).[22]

### 3.2.3. Control variables

Following prior literature, we add a number of risk factors as controls in our regression models (i.e. number of analysts' estimations, beta, leverage, book-to-price ratio, volatility of profitability, growth, and disclosure). These factors are standard controls in the cost of equity literature, which widely documents their association with measures of the cost of equity (e.g. Gebhardt et al., 2001; Gietzmann and Ireland, 2005; Francis et al., 2008).

*Number of estimates (NEST)*
In order to proxy for the level of attention received by a company we use the number of analysts' estimations of one-year-ahead EPS. This is a standard control variable used in disclosure-related studies. Starting from the seminal work by Botosan (1997), the previous literature has documented a strong influence of the level of analysts' attention on the relationship between disclosure and cost of equity capital.

*Beta (BETA)*
The capital asset pricing model (CAPM) predicts a positive association between the market beta of a stock and its cost of capital. However, previous studies do not consistently show such an expected relationship. While Botosan (1997) or Hail (2002) confirm the expected positive sign, Gebhardt et al. (2001) observe the expected sign but beta loses its significance when they add their industry measure. Finally, Francis et al. (2005) observe a negative relationship between beta and their measure of the cost of equity. We obtain the beta of each stock using a market model for the 60 months prior to

[21] As we are interested in obtaining an indicator of the level of distortion, absolute values are used in order to avoid the offsetting of positive and negative values of the individual RGD's.

[22] The direction of the results does not change when we use a GDI of 10% as the cut-off point. This level of distortion was found to affect users' perceptions in the study by Beattie and Jones (2002).

June of t+1, requiring, at least, 12 monthly return observations.

*Leverage (LEV)*
We measure leverage as the ratio between long-term debt and the market value of equity at 31 December of year *t*. Modigliani and Miller (1958) predict that the cost of equity should be increasing in the amount of debt in the financial structure of the company. This prediction is supported by results of studies such as those by Gebhardt et al. (2001) or Botosan and Plumlee (2005). In line with previous literature we expect to find a positive relationship between leverage and the cost of equity capital.

*Book-to-price ratio (BP)*
This is the ratio between book value of equity and market value of equity at 31 December of year *t*. Prior research documents a positive association between the book-to-price ratio and average realised returns (e.g. Fama and French, 1993 and Davis et al., 2000) as well as different ex-ante measures of the cost of equity (e.g. Gebhardt et al., 2001 and Botosan and Plumlee, 2005). These results are interpreted as evidence that the book-to-price ratio proxies for risk.

*Volatility of profitability (V_NI)*
Based on practitioners' consideration of the variability of earnings, Gebhardt et al. (2001) argue that this variable can be regarded as a source of risk. In our study, the volatility of the profitability of the firm is calculated as the standard deviation of net income scaled by mean of net income over a period of five years ending at December of year t.

*Growth (GROWTH)*
Following Francis et al. (2005) we control for the recent growth experienced by the company measured as the log of 1 plus the percentage change in the book value of equity along year t.

*Disclosure (RINDEX)*
Finally, to test whether there is an interaction between graph distortion and disclosure (H2), we introduce a measure of overall corporate disclosure. The relationship between corporate disclosure and the cost of equity is widely documented in the literature (e.g. Botosan and Plumlee, 2002; Espinosa and Trombetta, 2007; Francis et al., 2008). As an indicator of the information provided by the company we use the disclosure index published annually by the business magazine *Actualidad Económica*. This index is based on the information disclosed by companies in their annual reports. These reports are reviewed by the panel of experts who assign a score to a list of information items. For each company, the scores for each item are then added up to obtain a global score intended to represent the disclosure policy of the entity. Finally, the disclosure index is calculated as the ratio between the actual score of the company and the maximum possible score. Similar to Botosan and Plumlee (2002) or Nikolaev and Van Lent (2005), we use fractional ranks of the annual report indexes. Firms are ranked from 1 to N for each year and then the rank of each firm is divided by the total number of firms in this year to obtain the fractional ranks.

Table 2 provides a summary of the definition and data source of variables used in our analyses.

## 4. Results

### 4.1. Descriptive statistics

Descriptive statistics are presented in Table 3, where it can be seen that companies in our sample make wide use of graphs in their annual reports. At least one graph is included by 92% of the companies and the mean number of graphs per annual report (GRAPH) is 16. These figures are similar to those observed for other countries (e.g. Beattie and Jones, 2001). Although not reported in Table 3, mean RGDFAV is higher than 2.5% for 38% of the companies included in our study; that is, more than one third of the companies in our sample have favourably distorted the graphs included in their annual reports above the cut-off point chosen as an indication of material distortion. Unfavourable distortions are less frequent; mean RGDUNF is higher than 2.5% for 11% of the companies in our sample.

The correlation matrix presented in Table 4 shows that our ex-ante measure of the cost of equity is significantly correlated with all the risk proxies included in our study. As expected, the cost of equity is positively related to beta, book-to-price ratio, leverage and earnings variability and negatively related to growth and the number of estimates that acts also as a proxy for corporate size. Table 4 also shows a significant negative correlation between the cost of equity and our indicators of graph distortion.

### 4.2. Multivariate analysis

#### 4.2.1. Ex-ante measure of the cost of equity ($R_{PEG\_PREM}$)

This section presents the results obtained for the ex-ante measure of the cost of equity (i.e. the RPEG measure as developed by Easton, 2004). To asses the validity of this measure, we start our analysis by estimating a model similar to that developed by Botosan and Plumlee (2005) and used to test the relation between the cost of capital and a number of indicators of firm risk. We estimate the following equation:

$$R_{PREM\,it+1} = \alpha_i + \beta_1\,NEST_{it} + \beta_2 LEV_{it} + \beta_3\,V\_NI_{it} + \beta_4\,BETA_{it} + \beta_5\,BP_{it} + \beta_6\,GROWTH_{it} + e_{it}, \qquad (1)$$

**Table 2**
**Definition of variables**

| *Variables* | *Definitions and data source* |
|---|---|
| **Dependent variables** | |
| $R_{PEG\_PREM}$ | Estimated risk premium calculated as the $R_{PEG}$ in Easton (2004) less the risk-free rate (source JCF database). |
| $R_{PEF\_PREM}$ | Estimated risk premium calculated as the $R_{PEF}$ in Easton and Monahan (2005) less the risk-free rate (source JCF database). |
| $R_{MPEG\_PREM}$ | Estimated risk premium calculated as the $R_{MPEG}$ in Easton (2004) less the risk-free rate (source JCF database). |
| $R_{AVRG\_PREM}$ | Estimated risk premium calculated as the average of $R_{PEF}$, $R_{PEG}$, and $R_{MPEG}$ less the risk-free rate (source JCF database). |
| $R_{jm} - R_{fm}$ | Monthly excess return for firm *j* (source JCF database). |
| **Independent variables** | |
| GRAPH | Total number of graphs in the annual report of year t (source corporate annual reports). |
| RRGDFAV | Fractional rank of RGDFAV. RGDFAV is mean favourable Relative Graph Discrepancy (RGD) index across all graphs in the annual report. This index measures graph distortion and (0) means no distortion or distortion that is unfavourable to the company (source corporate annual reports). |
| RRGDUNF | Fractional rank of RGDUNF. RGDUNF is mean unfavourable Relative Graph Discrepancy (RGD) index across all graphs in the annual report. This index measures graph distortion and (0) means no distortion or distortion that is favourable to the company (source corporate annual reports). |
| RRGDFFAV | Fractional rank of RGDFFAV. RGDFFAV is mean favourable RGD of financial graphs (source corporate annual reports). |
| RRGDRFAV | Fractional rank of RGDRFAV. RGDRFAV is mean unfavourable RGD of non-financial graphs (source corporate annual reports). |
| $R_{Mm} - R_{fm}$ | Monthly excess return on the market portfolio (source JCF database). |
| $\text{RGDFAVfactor}_m$ | Return to the graph distortion factor mimicking portfolio (source JCF database). |
| **Control variables** | |
| NEST | Number of analysts' estimations of one-year-ahead EPS (source JCF database). |
| BETA | Capital market beta estimated via the market model with a minimum of 12 monthly returns over the 60 months prior to June of year t+1 (source JCF database). |
| BP | Book value of equity scaled by market value of equity (source OSIRIS database). |
| LEV | Long-term debt scaled by market value of equity (source OSIRIS database). |
| V_NI | Standard deviation of net income scaled by mean net income over a period of five years ending at December of year t (source OSIRIS database) . |
| GROWTH | Log of 1 plus the percentage change in the book value of equity along year t (source OSIRIS database). |
| ROA | Return on total assets (source OSIRIS database). |
| CDA | Current discretionary accruals estimated by using the model developed by Dechow and Dichev (2002) as modified by McNichols (2002) (source OSIRIS database). |
| RINDEX | Fractional rank of the disclosure index prepared by the business magazine *Actualidad Económica* for year t (source business magazine *Actualidad Económica*). |

**Table 3**
**Descriptive statistics**

| | *Mean* | *Std. Dev.* | *Minimum* | *Maximum* | *25th perc.* | *50th perc.* | *75th perc.* |
|---|---|---|---|---|---|---|---|
| $R_{PEG\_PREM}$ | 5.972 | 4.448 | –0.693 | 25.031 | 3.134 | 5.128 | 7.682 |
| NEST | 15.927 | 9.103 | 1 | 40 | 9 | 16 | 21 |
| BETA | 1.125 | 0.549 | 0.034 | 4.594 | 0.773 | 1.038 | 1.348 |
| BP | 0.637 | 0.403 | 0.049 | 3.018 | 0.362 | 0.558 | 0.818 |
| LEV | 0.398 | 0.603 | 0 | 6.315 | 0.065 | 0.221 | 0.469 |
| V_NI | 0.579 | 1.646 | –2.670 | 19.930 | 0.180 | 0.320 | 0.558 |
| GROWTH | 0.103 | 0.235 | –1.769 | 1.515 | 0.013 | 0.079 | 0.148 |
| GRAPH | 16.087 | 15.461 | 0 | 84 | 4.25 | 12 | 23 |
| RGDFAV | 0.106 | 0.307 | 0 | 3.685 | 0 | 0 | 0.095 |
| RGDUNF | 0.012 | 0.062 | 0 | 0.878 | 0 | 0 | 0 |
| RGDFFAV | 0.108 | 0.326 | 0 | 3.685 | 0 | 0 | 0.068 |
| RGDRFAV | 0.043 | 0.151 | 0 | 1.5 | 0 | 0 | 0 |
| ROA | 6.273 | 5.820 | –21.190 | 28.100 | 2.680 | 5.145 | 8.150 |
| CDA | 0.021 | 0.065 | –0.318 | 0.238 | –0.011 | 0.017 | 0.048 |
| INDEX | 0.637 | 0.150 | 0.230 | 0.960 | 0.550 | 0.630 | 0.750 |

The sample consists of 259 firm-year observations for the period 1996–2002. The definition of variables is provided in Table 2.

where:

$R_{PREM}$ = Proxy for the equity risk premium = $R_{PEG} - R_f$

$R_f$ = the risk-free rate, proxied by the interest rate on five-year Spanish Treasury bills.

The rest of the variables are defined in Table 2.

We estimate all our models using the fixed effects technique. The importance of using proper panel data estimation techniques when dealing with financial pooled data has been stressed by Nikolaev and Van Lent (2005) and Petersen (2005). If a simple OLS method is used to compute the estimated coefficients, their significance is very likely to be overstated. A traditional way to correct for this problem is to estimate yearly regressions and then take the average of the estimated coefficients, evaluating the statistical significance of these estimates by using the Fama-Macbeth t-statistic. This corrects for cross-sectional dependence, but not for time dependence. Nikolaev and Van Lent (2005) show how important firm effects can be when studying cost of capital determinants for a pooled sample of companies. This is our reason for estimating our model by using the fixed effects technique.[23]

Results of estimating equation (1) are presented in Table 5, Panel B (Model 1). The coefficients of the number of estimates, leverage, and variability of profitability have the expected sign and are statistically significant. The fact that the coefficients of beta and the book-to-price ratio are not statistically significantly different from zero could be due to the use of the fixed effect estimation technique. Beta and the book-to-price are risk factors that are specific for each company. Given that with the adopted estimation technique a specific intercept is estimated for each company, it is highly likely that the effect of these variables is already captured by these constants. Overall, our results confirm those already obtained for the Spanish market by Espinosa and Trombetta (2007) and support the validity of the measure of the cost of equity capital and the choice of the control variables.

We now move on to the main part of our empirical study and insert in our specification the indicators of mean favourable and unfavourable graph distortion across all graphs included in the annual report. Specifically, we now estimate the following equation (Model 2):

$$\begin{aligned} R_{PREM\,it+1} &= \alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} \\ &+ \beta_3\, V\_NI_{it} + \beta_4\, BETA_{it} + \beta_5\, BP_{it} \\ &+ \beta_6\, GROWTH_{it} + \beta_7\, RRGDFAV_{it} \\ &+ \beta_8\, RRGDUNF_{it} + \varepsilon_{it} \end{aligned} \qquad (2)$$

All variables are defined in Table 2.

We now find that favourable graph distortion is significantly and negatively related to the ex-ante measure of the cost of equity. A negative coefficient is also observed for unfavourable graph distortion but results show that it is insignificantly different from zero. These results suggest that favourably distorted graphs introduce a bias on users' perceptions. According to our results, annual report users perceive a better image of corporate performance when the annual report includes

[23] However, at least for our main analysis, we also provide the results obtained by running an OLS regression on the pooled sample. These results can be found in Table 4, Panel A.

**Table 4**
**Correlation matrix**

| | $R_{PEG\_PREM}$ | NEST | BETA | BP | LEV | V_NI | GROWTH | RGDFAV | RGDUNF | RGDFFAV | RGDRFAV | ROA | CDA |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| NEST | –0.200 | | | | | | | | | | | | |
| | (0.001) | | | | | | | | | | | | |
| BETA | 0.222 | –0.124 | | | | | | | | | | | |
| | (0.000) | (0.046) | | | | | | | | | | | |
| BP | 0.379 | –0.329 | 0.099 | | | | | | | | | | |
| | (0.000) | (0.000) | (0.112) | | | | | | | | | | |
| LEV | 0.172 | 0.161 | 0.023 | 0.417 | | | | | | | | | |
| | (0.006) | (0.009) | (0.707) | (0.000) | | | | | | | | | |
| V_NI | 0.192 | –0.121 | 0.168 | 0.091 | –0.020 | | | | | | | | |
| | (0.002) | (0.053) | (0.007) | (0.144) | (0.745) | | | | | | | | |
| GROWTH | –0.165 | 0.104 | –0.011 | –0.193 | –0.028 | 0.088 | | | | | | | |
| | (0.008) | (0.093) | (0.854) | (0.002) | (0.651) | (0.159) | | | | | | | |
| RGDFAV | –0.175 | 0.172 | 0.031 | –0.036 | 0.081 | –0.096 | 0.014 | | | | | | |
| | (0.005) | (0.005) | (0.616) | (0.559) | (0.195) | (0.122) | (0.822) | | | | | | |
| RGDUNF | –0.077 | 0.180 | –0.014 | –0.073 | 0.095 | –0.041 | 0.066 | 0.339 | | | | | |
| | (0.217) | (0.004) | (0.825) | (0.243) | (0.127) | (0.516) | (0.291) | (0.000) | | | | | |
| RGDFFAV | –0.202 | 0.112 | 0.025 | –0.048 | 0.034 | –0.086 | 0.089 | 0.872 | 0.276 | | | | |
| | (0.001) | (0.073) | (0.689) | (0.444) | (0.590) | (0.168) | (0.152) | (0.000) | (0.000) | | | | |
| RGDRFAV | –0.102 | 0.268 | –0.004 | –0.024 | 0.131 | –0.057 | –0.101 | 0.597 | 0.311 | 0.301 | | | |
| | (0.103) | (0.000) | (0.955) | (0.698) | (0.036) | (0.357) | (0.106) | (0.000) | (0.000) | (0.000) | | | |
| ROA | –0.250 | 0.065 | –0.153 | –0.343 | –0.543 | –0.012 | 0.213 | –0.046 | –0.111 | –0.010 | –0.054 | | |
| | (0.000) | (0.298) | (0.014) | (0.000) | (0.000) | (0.853) | (0.001) | (0.456) | (0.073) | (0.871) | (0.387) | | |
| CDA | 0.093 | –0.088 | 0.007 | 0.062 | 0.076 | –0.081 | 0.002 | 0.131 | –0.021 | 0.090 | 0.071 | 0.166 | |
| | (0.163) | (0.191) | (0.920) | (0.356) | (0.259) | (0.225) | (0.976) | (0.050) | (0.754) | (0.177) | (0.288) | (0.013) | |
| RINDEX | –0.220 | 0.466 | –0.003 | –0.064 | 0.247 | –0.055 | 0.010 | 0.212 | 0.238 | 0.170 | 0.283 | –0.323 | –0.143 |
| | (0.000) | (0.000) | (0.959) | (0.304) | (0.000) | (0.375) | (0.876) | (0.001) | (0.000) | (0.006) | (0.000) | (0.000) | (0.032) |

Table reports Spearman correlations. Significance levels are shown in brackets. The definition of variables is provided in Table 2.

**Table 5**
**Regression of the ex-ante measure of cost of equity ($R_{PEG\_PREM}$) on risk proxies and graph distortion**

Model 1: $R_{PEG\_PREM\ it+1} = \alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} + \beta_3 V\_NI_{it} + \beta_4 BETA_{it} + \beta_5 BP_{it} + \beta_6 GROWTH_{it} + \varepsilon_{it}$

Model 2: $R_{PEG\_PREM\ it+1} = \alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} + \beta_3 V\_NI_{it} + \beta_4 BETA_{it} + \beta_5 BP_{it} + \beta_6 GROWTH_{it} + \beta_7 RRGDFAV_{it} + \beta_8 RRGDUNF_{it} + \varepsilon_{it}$

Model 3: $R_{PEG\_PREM\ it+1} = \alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} + \beta_3 V\_NI_{it} + \beta_4 BETA_{it} + \beta_5 BP_{it} + \beta_6 GROWTH_{it} + \beta_9 RRGDFFAV_{it} + \beta_{10} RRGDRFAV_{it} + \varepsilon_{it}$

**Panel A: Pooled OLS regression**

| | *Model 1* | | *Model 2* | | *Model 3* | |
|---|---|---|---|---|---|---|
| *Variable* | *Coeff.* | *P-val* | *Coeff.* | *P-val* | *Coeff.* | *P-val* |
| Intercept | 4.382 | 0.000 | 4.936 | 0.000 | 5.101 | 0.000 |
| NEST | –0.059 | 0.039 | –0.051 | 0.091 | –0.057 | 0.064 |
| LEV | 1.149 | 0.128 | 1.122 | 0.145 | 1.136 | 0.143 |
| V_NI | 0.096 | 0.325 | 0.075 | 0.435 | 0.083 | 0.403 |
| BETA | 0.791 | 0.149 | 0.892 | 0.105 | 0.895 | 0.096 |
| BP | 2.089 | 0.010 | 2.111 | 0.010 | 2.053 | 0.011 |
| GROWTH | –2.172 | 0.087 | –2.131 | 0.097 | –1.839 | 0.160 |
| RRGDFAV | | | –1.721 | 0.072 | | |
| RRGDUNF | | | 0.162 | 0.900 | | |
| RRGDFFAV | | | | | –2.604 | 0.005 |
| RRGDRFAV | | | | | 0.910 | 0.530 |
| Adj. $R^2$ | 0.135 | | 0.138 | | 0.149 | |

**Panel B. Fixed effects regressions**

| | *Model 1* | | *Model 2* | | *Model 3* | |
|---|---|---|---|---|---|---|
| *Variable* | *Coeff.* | *P-val* | *Coeff.* | *P-val* | *Coeff.* | *P-val* |
| NEST | –0.147 | 0.011 | –0.130 | 0.027 | –0.126 | 0.032 |
| LEV | 2.537 | 0.004 | 2.491 | 0.005 | 2.516 | 0.005 |
| V_NI | 0.264 | 0.022 | 0.270 | 0.021 | 0.279 | 0.018 |
| BETA | –0.378 | 0.453 | –0.271 | 0.593 | –0.293 | 0.567 |
| BP | –0.109 | 0.912 | –0.127 | 0.897 | –0.182 | 0.851 |
| GROWTH | 0.160 | 0.867 | 0.208 | 0.831 | 0.296 | 0.766 |
| RRGDFAV | | | –2.351 | 0.013 | | |
| RRGDUNF | | | –0.182 | 0.828 | | |
| RRGDFFAV | | | | | –2.230 | 0.015 |
| RRGDRFAV | | | | | –1.032 | 0.260 |
| Adj. $R^2$ | 0.547 | | 0.554 | | 0.555 | |

The sample consists of 259 firm-year observations for the period 1996–2002. The definition of variables is provided in Table 2. Estimates of the firm-specific constant terms are omitted for readability.

graphs that have been distorted to portray a more favourable view of the company. The same is true if we focus our attention only on financial graphs (RRGDFFAV), as we do in Model 3. These results support our hypothesis H1. The information represented in graphs is usually included in the financial statements or other parts of the annual report. However, users get a different picture of the company when the annual report includes favourably distorted graphs. These results are consistent with the experimental evidence of the impact of improperly constructed graphs on subjects' choices provided by Arunachalam et al. (2002). They observe that students' decisions are affected by graph design. We extend these results by showing that, in a real setting, experts' (analysts') forecasts are biased because of distorted graphs included in annual reports.

We test the robustness of these results to the choice of the ex-ante measure of the cost of equity and to the inclusion of additional control variables. As for the proxy for the cost of equity, we calculate two alternative measures: $R_{PEF}$ and $R_{MPEG}$. The definition of these measures is given in the appen-

**Table 6**
**Regression of the cost of equity on risk proxies and graph distortion using fixed effects**

$$R_{PREM\,it+1} = \alpha i + \beta_1 NEST_{it} + \beta_2 LEV_{it} + \beta_3\, V\_NI_{it} + \beta_4\, BETA_{it} + \beta_5\, BP_{it} + \beta_6\, GROWTH_{it} + \beta_7\, RRGDFAV_{it} + \beta_9\, ROA_{it} + \beta_{10} CDA_{it} + \varepsilon_{it}$$

| | Panel A $R_{PEG\_PREM}$ | | Panel B $R_{PEF\_PREM}$ | | Panel C $R_{MPEG\_PREM}$ | | Panel D $R_{AVRG\_PREM}$ | |
|---|---|---|---|---|---|---|---|---|
| *Variable* | *Coeff.* | *P-val* | *Coeff.* | *P-val* | *Coeff.* | *P-val* | *Coeff.* | *P-val* |
| NEST | –0.121 | 0.029 | –0.177 | 0.000 | –0.085 | 0.111 | –0.125 | 0.003 |
| LEV | 1.590 | 0.075 | 1.032 | 0.071 | 1.847 | 0.046 | 1.410 | 0.021 |
| V_NI | 0.168 | 0.280 | –0.020 | 0.800 | 0.082 | 0.578 | 0.094 | 0.314 |
| BETA | –0.454 | 0.350 | –1.191 | 0.012 | –0.689 | 0.193 | –0.892 | 0.049 |
| BP | 0.510 | 0.701 | 0.971 | 0.253 | 0.774 | 0.573 | 0.692 | 0.451 |
| GROWTH | 1.082 | 0.287 | 0.748 | 0.299 | 1.150 | 0.283 | 0.726 | 0.307 |
| RRGDFAV | –2.232 | 0.012 | –1.327 | 0.041 | –2.645 | 0.002 | –2.296 | 0.001 |
| ROA | –0.334 | 0.000 | –0.037 | 0.378 | –0.327 | 0.000 | –0.172 | 0.000 |
| CDA | 3.090 | 0.277 | 5.413 | 0.010 | 5.793 | 0.049 | 5.296 | 0.014 |
| Adj. $R^2$ | 0.630 | | 0.599 | | 0.653 | | 0.659 | |

The sample consists of 259 firm-year observations for the period 1996–2002. The definition of variables is provided in Table 2. Estimates of the firm-specific constant terms are omitted for readability.

dix. We also use the average ($R_{AVRG}$) of the three proxies calculated in our study. Additionally, we check if our results are driven by factors such as corporate performance or accruals quality. Prior literature documents a positive association between corporate performance and graph distortion (e.g. Beattie and Jones, 1999, 2000). Therefore, the effect of graph distortion on the cost of equity that we observe in this study could be driven by the fact that distorting companies are also those with the highest performance. Hence, we add a measure of corporate performance (i.e. ROA) as a control variable. Second, Francis et al. (2005) and Francis et al. (2008) show the existence of a positive relationship between the cost of equity capital and the (poor) quality of accruals. The negative relationship between graph distortion and the cost of equity observed in this study could be reflecting this association. Therefore, we test for the robustness of our results by adding a measure of accruals quality (i.e. discretionary accruals) to our model (CDA).[24]

Table 6 reports the results of these sensitivity analyses and shows that the effect of favourable graph distortion on the ex-ante cost of equity is robust to the choice of the measure of the cost of equity. Favourable graph distortion is found to be negatively and significantly related to all the alternative measures of the ex-ante cost of equity calculated in our study. Additionally, we observe that, although corporate performance and accruals quality can be highly significant variables in explaining the ex-ante cost of equity, its inclusion in the model does not qualitatively change our results. That is, the effect of favourable graph distortion on the ex-ante measures of the cost of equity remains.

*Interaction analysis*

To investigate whether the relationship between graph distortion and the ex-ante measure of the cost of equity varies depending on the overall level of disclosure, as stated in our second hypothesis, we introduce a measure of the voluntary information provided by the company in their annual report (RINDEX) and an interaction term between disclosure and graph distortion. Specifically, we estimate the following equation:

$$\begin{aligned} R_{PREM\,it+1} &= \alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} \\ &+ \beta_3\, V\_NI_{it} + \beta_4\, BETA_{it} + \beta_5\, BP_{it} + \beta_6\, GROWTH_{it} \\ &+ \beta_7\, RRGDFAV_{it} + \beta_8\, RINDEX_{it} \\ &+ \beta_9\, RRGDFAV_{it} * RINDEX_{it} + \varepsilon_{it} \end{aligned} \tag{3}$$

All variables are defined in Table 2.

Table 7, Panel A presents the results of the estimation of Equation (3). Consistent with the results reported previously, the coefficient on RRGDFAV is found to be negative and significant. A negative relationship is also observed between the overall level of disclosure and the cost of equity, although it is not significant at conventional levels. Finally, the interaction term is positive and significant, which means that graph distortion and disclosure interact in shaping their effects on the cost of equity. Since the sign of the interaction term is posi-

[24] To obtain this measure we use the model developed by Dechow and Dichev (2002) as modified by McNichols (2002).

**Table 7**
**Regression of the cost of equity on risk proxies, graph distortion, and disclosure using fixed effects**

Model 1: $R_{PEG\_PREM\,it+1}$ (Panel A) = $\alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} + \beta_3\, V\_NI_{it} + \beta_4\, BETA_{it} + \beta_5\, BP_{it} + \beta_6\, GROWTH_{it} + \beta_7\, RRGDFAV_{it} + \beta_8\, RINDEX_{it} + \beta_9\, RRGDFAV_{it} * RINDEX_{it} + \varepsilon_{it}$

Model 2: $R_{PEG\_PREM\,it+1}$ (Panels B–C) = $\alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} + \beta_3\, V\_NI_{it} + \beta_4\, BETA_{it} + \beta_5\, BP_{it} + \beta_6\, GROWTH_{it} + \beta_7\, RRGDFAV_{it} + \beta_8\, RINDEX\_D_{it} + \beta_9\, RRGDFAV_{it} * RINDEX\_D_{it} + \varepsilon_{it}$

| | Panel A | | Panel B | | Panel C | |
|---|---|---|---|---|---|---|
| *Variable* | *Coeff.* | *P-val* | *Coeff.* | *P-val* | *Coeff.* | *P-val* |
| NEST | –0.148 | 0.032 | –0.148 | 0.028 | –0.153 | 0.022 |
| LEV | 2.433 | 0.001 | 2.456 | 0.001 | 2.495 | 0.001 |
| V_NI | 0.284 | 0.050 | 0.266 | 0.066 | 0.276 | 0.058 |
| BETA | –0.350 | 0.555 | –0.345 | 0.564 | –0.161 | 0.786 |
| BP | –0.202 | 0.829 | –0.147 | 0.876 | –0.135 | 0.885 |
| GROWTH | 0.181 | 0.856 | 0.111 | 0.911 | 0.237 | 0.809 |
| RRGDFAV | –6.227 | 0.002 | –4.198 | 0.003 | –3.909 | 0.002 |
| RINDEX | –4.204 | 0.099 | | | | |
| RRGDFAV*RINDEX | 8.106 | 0.024 | | | | |
| RINDEX_D | | | –1.517 | 0.178 | –2.093 | 0.143 |
| RRGDFAV*RINDEX_D | | | 3.785 | 0.058 | 4.871 | 0.028 |
| Adj. $R^2$ | 0.564 | | 0.561 | | 0.565 | |

The sample consists of 259 firm-year observations for the period 1996–2002. The definition of variables is provided in Table 2. Estimates of the firm-specific constant terms are omitted for readability.

tive, results indicate that the effect of graph distortion on the cost of equity is moderated by the level of overall disclosure. Stated in other words, the effect of graph distortion on the cost of equity is partially (or eventually completely) offset by its interaction with the overall level of disclosure. To get a clearer picture of the interaction between graph distortion and disclosure we dichotomise the RINDEX variable and estimate the following equation:

$$R_{PREM\,it+1} = \alpha_i + \beta_1 NEST_{it} + \beta_2 LEV_{it} + \beta_3\, V\_NI_{it} + \beta_4\, BETA_{it} + \beta_5\, BP_{it} + \beta_6\, GROWTH_{it} + \beta_7\, RRGDFAV_{it} + \beta_8\, RINDEX\_D_{it} + \beta_9\, RRGDFAV_{it} * RINDEX\_D_{it} + \varepsilon_{it} \quad (4)$$

where all the variables are defined as before except for:

*RINDEX_D* = A dichotomy variable which, in Panel B (C), takes the value of one (1) when the value of the RINDEX is in the top 50% (33%) percentile of the distribution of this variable, and a value of zero (0) otherwise.

Results of this estimation are presented in Table 7, Panels B and C. Since the selection of the cut-off point used to dichotomise the RINDEX variable is arbitrary, Table 7 reports the results obtained by using two different cut-off points. In panel B, the set of high disclosers comprises those companies the disclosure index of which is above the median, while in Panel C high disclosers are those firms falling in the top 33% percentile of the distribution of the RINDEX variable.

Consistent with the results presented in Panel A, we find that the cost of equity is negatively associated with RRGDFAV and positively related to the interaction term. The dichotomisation of the variable RINDEX facilitates the interpretation of the results. When RINDEX_D takes the value of zero (0), the relationship between graph distortion and the cost of equity is given by $\beta_7$ and is negative, both in Panels B and C. This means that for low disclosers, graph distortion is negatively related to the cost of equity. However, when RINDEX_D takes the value of one (1), the effect of graph distortion is given by the addition of the coefficients on RRGDFAV and the interaction term (i.e. $\beta_7 + \beta_9$). This addition results in a negative figure (–0.413) in Panel B, which is much lower than the coefficient on RRGDFAV, and a positive figure (0.962) in Panel C. Thus, for transparent companies, disclosure partially removes the effect of graph distortion (Panel B) or even transforms it into a positive effect (Panel C). These results indicate that disclosure moderates the relationship between graph distortion and the cost of equity and provide support for our second hypothesis. Furthermore, differences between Panels B and C indicate that the moderating effect is increasing in

**Table 8**
**Firm-specific regressions of stock returns on the market portfolio and the graph distortion factor**

Model 1: $R_{jm} - R_{fm} = a_j + b_j (R_{Mm} - R_{fm}) + \varepsilon_{jm}$
Model 2: $R_{jm} - R_{fm} = a_j + b_j (R_{Mm} - R_{fm}) + c_j$ RGDFAVfactor$_m + \varepsilon_{jm}$

| | Model 1 | | Model 2 | |
|---|---|---|---|---|
| *Variable* | *Coeff.* | *P-val* | *Coeff.* | *P-val* |
| Constant | 0.004 | 0.529 | 0.005 | 0.520 |
| $R_{Mm} - R_{fm}$ | 0.836 | 0.020 | 0.851 | 0.012 |
| RGDFAVfactor$_m$ | | | −0.130 | 0.326 |
| R-squared | 0.264 | | 0.296 | |

The table reports the average coefficient estimates obtained from the estimation of the asset-pricing models for each company included in our sample. A minimum of 18 monthly stock returns for the period June 1997 to May 2004 is required. The definition of variables is provided in Table 2.

the level of disclosure, so that when transparency is sufficiently high, the relationship between graph distortion and the ex-ante measure of the cost of equity becomes positive. Results suggest that while at low levels of transparency graph distortion remains undetected, high levels of disclosure uncover graph distortion, which results in an increase in the risk perceived by users. Hence, users' predictions can only be potentially affected by graph distortion when the level of information provided by the company is low.

So far we have used an ex-ante measure of the cost of equity based on analysts' estimations of earnings per share. Hence, our results could be interpreted as evidence that graph distortion affects users' (analysts') perceptions of corporate performance. However, it is important to check that the market is able to correct individuals' biases. With this aim, in the next section we present the results obtained using average ex-post returns as a proxy for the cost of equity.

*4.2.2. Ex-post measure of the cost of equity (stock returns)*

As a starting point in our ex-post analysis, we estimate a one-factor asset-pricing model for each of the 68 companies in our sample.[25] The average coefficients and adjusted R's squared of these estimations are presented in Table 8 (Model 1).

Results show a mean beta of 0.84 and a mean adjusted R-squared of 26.4%. We proceed by adding a factor aimed at representing favourable graph distortion (RGDFAVfactor). Specifically, we estimate the following equation:

$$R_{jm} - R_{fm} = a_j + b_j (R_{Mm} - R_{fm}) + c_j \, \text{RGDFAVfactor}_m + \varepsilon_{jm} \quad (5)$$

[25] A minimum of 18 monthly returns is required in these estimations.

where:

$R_{jm} - R_{fm}$ = monthly excess return for firm j.

$R_{Mm} - R_{fm}$ = monthly excess return on the market portfolio

RGDFAVfactor$_m$ = return to the graph distortion factor mimicking portfolio

Average coefficients obtained from firm-specific estimations of Equation (5) are reported in Table 8 (Model 2). Results show that stock returns are not affected by the distortion of graphs included in the annual report. The coefficient of the RGDFAV factor is negative but it is insignificantly different from zero. Hence, we find support for our hypothesis H3. Results (unreported) are similar when we construct the RGDFAV factor relaying on favourably distorted graphs representing financial variables. Although these variables might exert a higher influence on users' decisions than non-financial variables, results show that stock returns are not affected by the distortion of financial graphs. These results are in accordance with the EMH and suggest that decision makers (at the aggregated level) are able to see through distortion. Their decisions cannot be biased by means of 'rosy' graphs depicting a much more favourable view of corporate performance than that reflected in the financial statements. Nonetheless, results presented in the previous section show the existence of a negative and significant relationship between favourable graph distortion and the ex-ante measure of the cost of equity. One way to explain these apparently contradicting findings is that individuals' perceptions can be biased because of distorted graphs, but the aggregation process performed by the stock market leads to unbiased decisions. Prior research shows that the aggregation of individual's predictions leads to higher levels of accuracy (e.g. Solomon, 1982 and Chalos, 1985). Furthermore,

Easton and Sommers (2007) provide evidence showing that the market undoes the optimistic bias in analysts' forecasts. Therefore, the possibility exists that analysts' estimations of earnings per share (the base for the ex-ante measure of the cost of equity) are influenced by graph distortion, even though stock prices resulting from the aggregation of a large number of decisions are not biased by distorted graphs.

To further investigate this issue, following Easton and Sommers (2007), we recalculate the $R_{PEG}$ measure using realised earnings instead of analysts' earnings forecasts. By doing so the $R_{PEG}$ measure becomes an ex-ante measure of the cost of equity based on perfect foresight, since it is not based on estimations but on realised earnings. We re-estimate Equation (2) using this measure as the dependent variable. Unreported results, available from the authors, show that this measure of the cost of equity is not affected by graph distortion. Even when we focus our attention on financial graphs we do not observe any significant relationship between graph distortion and this perfect foresight measure of the cost of equity. These results corroborate the intuition gathered from previous analyses. Individuals' (analysts') perceptions are biased and distorted graphs affect this bias, but the market is able to correct these misperceptions. Once this bias in expectations is removed, the effect of graph distortion on the cost of equity disappears. Our results are consistent with those reported in previous experimental studies showing that individuals' perceptions are affected by graph distortion (e.g. Beattie and Jones, 2002 and Arunachalam et al., 2002). We extend these results by providing evidence of the effect of graph distortion on individuals' (analysts') estimations in a real setting, instead of a controlled experiment. Additionally, we show that the bias in individuals' expectations is corrected in the capital market. Hence, despite the efforts made by companies to create a more favourable impression on users, we provide evidence showing that the aggregated market response is not biased by distorted graphs.

## 5. Conclusions

The wide use of graphs in corporate annual reports together with the frequency with which they are distorted to portray a more favourable view of corporate performance, suggest that companies expect benefits from using and distorting graphs. In other words, distorted graphs are used by companies as an impression management technique. This 'manipulation hypothesis' is confirmed, for example, by the time-series analysis of Beattie and Jones (2000). However, the information depicted in graphs is usually presented also in a numerical format in the annual report. Using the terminology of Merkl-Davies and Brennan (2007), graphs do not provide any 'incremental information'. Consequently, market efficiency would imply that stock prices should not be affected by the use of distorted graphs to display information already available in the financial statements or other sections of the annual report. We can call this second hypothesis the 'absence of incremental information' hypothesis. To the best of our knowledge, our study is the first to test these two hypotheses by investigating the effect of graph distortion on the cost of equity capital both on an ex-ante and an ex-post basis.

Using a sample of companies quoted on the MSE, we obtain the following results. In accordance with the EMH, we find that ex-post stock returns are not affected by the distortion of graphs included in the annual report. Nonetheless, we observe a robust negative relationship between favourable graph distortion and ex-ante measures of the cost of equity. This negative relationship, however, is moderated by the overall level of disclosure, so that at high levels of transparency the relationship between graph distortion and the ex-ante measure of the cost of equity becomes positive. Results suggest that disclosure uncovers graph distortion which in turn leads to a higher risk being perceived.

In order to understand our results, it is important to distinguish between analysts' expectations and aggregated market behaviour. Since ex-ante measures of the cost of equity are based on analysts' forecasts, results indicate that analysts' estimations can be biased when graphs are distorted to portray a more favourable view of the company, especially when the overall level of information provided is low. These results are consistent with existing theories on the effects of herding and limited attention on the use of information in capital markets.[26] These theories predict that a focus on relative performance evaluation can induce analysts to neglect some of their detailed private information and to focus on coarser and more readily available indicators that summarise the existing information.

After analysts' forecasts have been formed, investors take their buying and selling decisions and these decisions determine the ex-post measure of the cost of equity, which is not found to be affected by distorted graphs. Hence, the limited attention effects are only transitory, because they disappear once the market settles. This means that with the passing of time, market activity is able to correct individuals' misperceptions, so that stock returns are not affected by distorted graphs. These results are related to those reported by Easton and Sommers (2007) showing that the market undoes the optimistic bias in analysts' forecasts. However, we extend their analysis in two ways. First, we

[26] These theories have been reviewed in the previous sections.

highlight graph distortion as a significant determinant of this bias. Second, we show that their result is robust to the use of a three-factor model as a measure of ex-post cost of equity as proposed by Fama and French (1993).

Since the analysis was conducted in a relatively underdeveloped capital market, where herding behaviour and heuristic trading are more likely, we conjecture that, in spite of the documented prevalence of graph usage and graph distortion in countries with highly developed capital markets (e.g. the UK and the US), it is unlikely that firms in these countries will achieve a reduction in their realised cost of equity by means of distorted graphs. However, further research is needed to investigate whether analysts' estimations are also biased because of distorted graphs in highly developed capital markets.

Our empirical evidence also sheds some additional light on the relationship between disclosure and the cost of equity. Both Espinosa and Trombetta (2007) and Gietzmann and Ireland (2005) show the important role played by accounting policy choice as a conditioning factor in this relationship. Our results show that disclosure interacts also with graph distortion as a determinant of the cost of equity.

More generally, we provide some initial evidence on the economic effects of materially distorted graphs. The previous literature had already documented the fact that a considerable number of graphs presented in annual reports are distorted. The minimum level of distortion necessary to influence users' perceptions was also established through experimental studies. Building on these important findings, we have extended the analysis and investigated the effects of graph bias on the cost of equity. However, this is only one of the possible economic variables that can be affected. The task of investigating the effects of distorted graphs on other economic variables is left to future research.

Finally, Steinbart (1989) and Beattie and Jones (2002) advocate regulatory intervention and the involvement of auditors in order to eliminate graph distortion. This proposal is based on the presumption that any misrepresentation of firms' financial performance potentially misleads users of annual reports. Our results have some important implications for this debate. Drawing on prior experimental evidence, we enhance our understanding of the significant economic effects of graph distortion outside a controlled laboratory environment. The fact that, over time, the market is able to correct individuals' misperceptions caused by distorted graphs does not diminish their importance with regard to the continuous functioning of the stock market. Consequently, our results support the notion that there is a rationale for the enforcement of regulatory action to avoid significant graph distortion. At the same time, our findings also underscore the complexity of the influences that graphs have on economic decisions and they call for an increased effort to advance knowledge on how graphs affect economic behaviour.

## Appendix

*Measures of the ex-ante cost of equity capital*

All three alternative proxies for the ex-ante cost of equity capital that we use in the paper are based on the valuation model known as the Abnormal Earnings Growth Valuation Model. The general valuation formula according to this model is as follows:

$$P_t = \frac{x_{t+1}}{r} + \sum_{\tau=2}^{+\infty} \frac{x_{t+\tau} + rd_{t+\tau-1} - (1+r)x_{t+\tau-1}}{r} = \frac{x_{t+1}}{r} + \sum_{\tau=2}^{+\infty} \frac{AEG_{t+\tau}}{r(1+r)^{\tau}} \qquad \text{(A.1)}$$

where:

$P_t$ = closing price of the month of June
$x_t$ = earnings at time $t$
$r$ = cost of equity capital
$d_t$ = the dividend payout at time $t$
$AEG_t$ = abnormal earnings growth rate at time $t$.

Starting from this general valuation formula, each of the three proxies is obtained by making some restrictive assumptions on the parameters of the model (cf. Easton and Monahan, 2005). The expressions for each of the three measures are as follows:

a) Price to Forward Earnings Model (PEF)

$$P_0 = \frac{eps_1 + r_e d + eps_2}{(r_e + 1)^2 - 1} \qquad \text{(A.2)}$$

## Appendix *(continued)*

b) Price to Earnings Growth model (PEG)

$$P_0 = \frac{eps_2 - eps_1}{(r_e)^2} \quad \text{(A.3)}$$

c) Modified Price to Earnings Growth models (MPEG)

$$P_0 = \frac{eps_2 + r_e d - eps_1}{(r_e)^2} \quad \text{(A.4)}$$

$P_0$ = price of the stock of the company at 30 June of year $t$+1
$d$ = dividend payout ratio for year $t$
$eps_1$ = one year ahead consensus forecast of earnings per share at 30 June of year $t$+1
$eps_2$ = two years ahead consensus forecast of earnings per share at 30 June of year $t$+1
Solving for $r_e$ in each of these equations produces the estimates of the ex-ante cost of equity.

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# Value-relevance of presenting changes in fair value of investment properties in the income statement: evidence from Hong Kong

**Stella So and Malcolm Smith***

**Abstract**—This study investigates the value-relevance of the revision introduced in HKAS 40 (2004) 'Investment Property' on the presentation of changes in the fair value of investment properties. The revision follows that introduced in IAS 40 (2000) as Hong Kong adopted the International Financial Reporting Standards in 2005. As introduced in IAS 40 (2000), HKAS 40 (2004) requires that companies choosing to adopt the fair value model have to present changes in the fair value of investment properties in the income statement. Previously under the Hong Kong accounting standard SSAP 13 (2000), such changes were presented primarily in the revaluation reserve. Using both a three-day short window centred around the earnings announcement date and a 12-month long window, this study provides evidence that investors value the HKAS 40 (2004) revision in the presentation of the changes in fair value of investment properties. Based on a sample of listed property companies in Hong Kong during 2004–2006, the results of this study show a significantly higher market price reaction and returns association when changes in fair value of investment properties are presented in the income statement. The results of this study are of interest not only to academic researchers, but to practitioners and standard setters as they assess the decision usefulness of the revised presentation.
**Keywords:** financial reporting; Hong Kong; investment; property; fair value

## 1. Introduction

This study examines the value-relevance of the accounting information reported as a consequence of Hong Kong Accounting Standard No. 40 'Investment Property' (HKICPA, 2004), hereafter HKAS 40 (2004), on the presentation of changes in fair value of investment properties. The date 1 January 2005 marked the beginning of a new era when accounting standards in Hong Kong became fully converged with International Financial Reporting Standards (IFRS). As part of the final phase of the convergence, a number of new accounting standards which are word-for-word equivalents of the IFRSs were issued to replace the existing pronouncements. HKAS 40 (2004) is one of them. Following International Accounting Standards No. 40 'Investment Properties' (2000) (IASC, 2000), hereafter IAS 40 (2000),[1] HKAS 40 (2004) requires that companies choosing to adopt the fair value model must present changes in the fair value of investment properties in the income statement. Previously under Statement of Standard Accounting Practice No. 13 'Accounting for Investment Properties' (HKSA, 2000), hereafter SSAP 13 (2000), such changes (labelled, however, as open market value[2] changes) were presented

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[1] The fair value model discussed in this paper is introduced in IAS 40 (2000). Although it had been revised in 2003 and superseded by IAS 40 (2003) (IASCF, 2008a), the revision in 2003 is related to clarification of aspects of property interest that are held under an operating lease and is therefore of no relevance to this study. Of relevance to this study are the fair value model and the required presentation of fair value changes of investment properties in the income statement under the fair value model, which were introduced in IAS 40 (2000). Throughout this paper, reference is therefore made to IAS 40 (2000) and not to IAS 40 (2003).

[2] SSAP 13 (2000) adopted the definition of open market value from the Hong Kong Institute of Surveyors which defined it as the best price at which an interest in a property might reasonably be expected to be sold at the date of valuation assuming a willing seller, a reasonable period in which to negotiate the sale taking into account the nature of the property and the state of market, that values will remain static during that period, that the property will be freely exposed to the open market, and that no account will be taken of any additional bid by a purchaser with a special interest (HKSA, 2000). The IASB defines fair value as the amount for which an asset could be exchanged between knowledgeable, willing parties in an arm's length transaction (IASCF, 2008a). The former IASC, in its Basis for Conclusions on IAS 40 (2000), states that its concept of fair value for investment properties is similar to the International Valuation Standards Committee (IVSC) concept of market value. IVSC defines market value as the estimated amount for which an asset should exchange on the date of valuation between a willing buyer and a willing seller in an arm's length transaction after proper marketing wherein the parties had each acted knowledgeably, prudently and without compulsion (IASCF, 2008c).

primarily[3] in a different location, the revaluation reserve.

Changes in the fair value of investment properties, whether increases or decreases, represent gains or losses which are unrealised. There have been concerns that the inclusion and presentation of such unrealised gains and losses in the income statements might lead to undue increases in earnings volatility and investor confusion (HKSI, 2006). The former International Accounting Standards Committee (IASC) argues, however, in its Basis for Conclusions on IAS 40 (2000) that such presentation provides the most relevant and transparent view of the financial performance of investment properties (IASCF, 2008c). After all, the objective of financial statements is not to smooth profit figures (McBride, 2006), but to reflect economic reality for economic decision making (IASCF, 2008b).

The replacement of SSAP 13 (2000) by HKAS 40 (2004) in Hong Kong thus provides a unique opportunity for this study to examine the issues relating to the impact of the presentation location, whether in the income statement or revaluation reserve, for changes in fair value of investment properties. The results of this study will have implications for companies around the world that prepare their financial statements using International Financial Reporting Standards.

Like all other value-relevance studies, this study use share prices and share returns to infer whether investors consider accounting information to be sufficiently relevant and reliable to be useful in making investment decisions (Maines and Wahlen, 2006; Landsman, 2007). Value-relevance tests are generally joint tests of relevance and reliability[4] (Barth et al., 2001), where reliability is more than agreement about a measure (measurement verifiability), but also involves the correspondence between description, classification and presentation (representational faithfulness), of the phenomenon it purports to represent (IASCF, 2008b; Maines and Wahlen, 2006). This study examines the relevance and reliability of the HKAS 40 (2004) revision on the presentation of changes in fair value of investment properties, by evaluating the revision's ability to impact upon the abnormal share returns to investors.

There are three distinct issues which motivate this study.

First, while the efficient market hypothesis suggests that the presentation location of accounting information is not valued by investors, results from prior research have shown otherwise (e.g. Hirst and Hopkins, 1998; Maines and McDaniel, 2000; Barth et al., 2003; Hirst et al., 2004; Chambers et al., 2006; and Lee et al., 2006). The presentation location of an accounting amount affects not only its relevance but also its reliability (IASCF, 2008b; Maines and Wahlen, 2006). Given these prior findings and associated theory, it is surprising that Owusu-Ansah and Yeoh (2006) find no difference whether unrealised gains (losses were not studied) on investment properties in New Zealand are reported in the income statement or revaluation reserve. Further research is therefore needed to provide more evidence about the impact of presentation location generally, and of changes in fair value of investment properties in particular.

Second, existing research (see Landsman (2007) for a detailed summary) focuses on financial assets and liabilities (Barth, 1994; Barth et al., 1995, 1996, 2006); Eccher et al., 1996; Nelson, 1996, Carroll et al., 2003; Hirst et al., 2004; Hodder et al., 2006; Danbolt and Rees, 2008) and employee share options (Espahbodi et al., 2002; Robinson and Burton, 2004). While there are studies on non-financial assets, for intangible assets and tangible long-lived assets (e.g. Barth and Clinch, 1998; Aboody et al., 1999; Muller and Riedl, 2002) and on investment properties (i.e. Dietrich et al., 2001; Owusu-Ansah and Yeoh, 2006), the empirical evidence is largely based on fair value disclosures in the notes to the accounts from the 1990s. More studies should be conducted using data after the implementation of the fair value accounting standards (e.g. IAS 39 'Financial Instruments: Recognition and Measurement', and HKAS 40 'Investment Property' studied in this study).

[3] Under SSAP 13 (2000), investment properties were to be included in the balance sheet at their open market value and changes in the open market value were to be recognised primarily in the revaluation reserve, i.e. in the revaluation reserve unless (1) its balance was insufficient to cover a deficit, in which case the amount by which the deficit exceeded the revaluation reserve balance was to be charged to the income statement; and (2) a revaluation surplus subsequently arises, this surplus was to be credited to the income statement to the extent of the deficit previously charged (HKSA, 2000).

[4] This paper assumes the application of the Framework for the Preparation and Presentation of Financial Statements published by the former International Accounting Standards Committee (IASC) in 1989 and re-adopted by the International Accounting Standards Board (IASB) in 2001 (hereafter Framework (1989), IASCF, 2008b). In Framework (1989), relevance and reliability are two of the four principal qualitative characteristics of decision-useful financial statements (IASCF, 2008b). Information is relevant if it influences the economic decisions of users by helping them evaluate past, present or future events or confirming, or correcting, their past evaluations. Information is reliable when it is free from material error and bias and can be depended upon by users to represent faithfully that which it purports to represent or could reasonably be expected to represent. Of note is that the term reliability in Framework (1989) has been proposed to be replaced by 'faithful representation' in the Exposure Draft of an Improved Conceptual Framework for Financial Reporting, jointly published in May 2008 by the IASB and the US Financial Accounting Standards Board (FASB). The boards considered that faithful representation encompasses all the key qualities that Framework (1989) includes as aspects of reliability and therefore proposed the replacement. No attempt is however made in this paper to anticipate the outcome of the exposure draft, which is expected to be finalised in the first half of 2009.

Third, Hong Kong is a world-recognised centre for property construction, development and investment (KPMG China and Hong Kong, 2004) and the total market capitalisation of its property industry (conglomerates/consolidated enterprises excluded) in 2006 was HK$13,249bn, representing 11% of the total market capitalisation of all Hong Kong Stock Exchange Main Board equities (HKSE, 2005, 2006). Investment property is a significant component of many company balance sheets in Hong Kong and the way it is accounted for has become an issue of prominent interest in Hong Kong in 2005–2006 (McBride, 2006).

This study employs a sample of listed companies from the Hong Kong property industry and examines whether the relevance and reliability of disclosures are enhanced through the adoption of HKAS 40 (2004) and the reporting of changes in fair value of investment properties in the income statement. Only companies from the property industry are included because these companies are expected to hold substantial levels of investment properties on their balance sheets and their earnings are thus likely to be more sensitive to the adoption of HKAS 40 (2004). Depending on the sample company's accounting year-end, and the incidence of early adoption of HKAS 40 (2004), the sample period in this study extends from 2004 to 2006.

Using models adapted from Easton and Harris (1991), Amir et al. (1993) and Barth (1994), results from this study provide evidence on the higher value-relevance of presenting fair value gains or loss in the income statement versus presenting them in the revaluation reserve. The higher value-relevance is found to be evident in both the short-window market reaction to the release of annual earnings announcements and the long-window market-adjusted annual returns. Taken together, our results give support to the HKAS 40 (2004) presentation which investors appear to value more than previously under SSAP 13 (2000).

The paper proceeds as follows. Section 2 describes related prior research and Section 3 details the research method employed. Section 4 discusses the empirical results, and the paper concludes in Section 5 with a summary and discussion of research opportunities.

## 2. Related prior studies

IAS 40 (2000) represents the first time that the IASB permits a fair value model for non-financial assets (IASCF, 2008c). Under the fair value model, investment properties are carried at fair values and changes in fair value, whether up or down, are included in the profit or loss for the period and presented in the income statements. Supporters of the fair value model believe that fair values give users of financial statements more useful information than other measures, such as depreciated cost, and changes in fair value are inextricably linked as integral components of the financial performance of an investment property and are therefore presented in the income statements (IASCF, 2008c). Although IAS 40 (2000) permits entities to choose between a fair value model or a cost model, the Basis for Conclusions on IAS 40 (2000) states clearly that it is highly unlikely that a subsequent change from the fair value model to the cost model can be made on the grounds of more appropriate presentation (IASCF, 2008c).

However, Penman (2007) does not entirely agree; he evaluates historical cost and fair value accounting from two perspectives – equity valuation and stewardship and concludes that while fair value accounting is a plus at a conceptual level, the minuses add up with fair value implemented as exit price (whether estimated or observed in active markets) and the problems with historical cost accounting remains unresolved.

Singleton-Green (2007) summarises the problems of fair value accounting as: (1) the lack of active markets for most assets and liabilities, which means that most fair value measurements are estimates and are highly subjective and potentially unreliable; (2) costly information, especially for smaller companies; and (3) the recognition of profits based on fair values, which mean that unrealised profits or losses from changes in fair value are recognised, and result in greater volatility and unpredictability. This study focuses on the third issue, the presentation of changes in fair value of investment properties, in the income statement versus the revaluation reserve.

Empirical studies assessing the relevance and reliability of fair value accounting versus historical cost-based accounting focus on financial instruments, and the results from these studies are generally mixed. Barth (1994) finds that, for a sample of US banks with data from 1971–1990, disclosed fair value estimates of investment securities provide significant incremental[5] explanatory power for bank share prices beyond that provided by historic costs. Fair value gains and losses of investment securities (constructed from two annually disclosed fair value estimates) are, however, found to have no significant incremental explanatory power for annual returns (changes in share price), due to the increased measurement errors (Barth, 1994). Similar results are obtained in Barth et al. (1995), Barth et al. (1996), Eccher et al. (1996) and Nelson (1996), all using bank data. Results from Carroll et al. (2003) differ; instead of using bank data, they sample closed-end mutual funds which typically have investment securities (report-

[5] According to Biddle et al. (1995), an incremental comparison determines whether one accounting measure provides information content beyond that provided by another.

ed at fair values) comprising virtually all their assets and with negligible liabilities and other assets. This is an advantage because the potential problem introduced by measuring some assets and liabilities at fair value but others at historical cost, is eliminated. Significant association between share prices and the fair value of investment securities, as well as between share returns and fair value securities gains and losses are found. To examine whether differences in the reliability of the fair value of investment securities affect investors' assessments of the usefulness of the information, Carroll et al. (2003) examine the association between share prices and fair values across different fund types and find that in all cases, including those traded in thin markets, there is a significant association between the share prices and fair values.

In contrast, Danbolt and Rees (2008), using UK data, report no support for full fair value accounting. While fair value income is considerably more value-relevant than historic cost income, the higher relevance disappears in the presence of changes in fair value accounting balance sheet values. Danbolt and Rees (2008) interpret their results as evidence of the absence of an obvious advantage from adopting fair value income accounting if fair value balance sheet values are available to the user.

Value-relevance research studies the association between fair value estimates and share prices or returns. Sloan (1999) comments that while this association provides evidence that investors find fair value estimates to be relevant, the inferences regarding reliability are indirect and limited by the fact that share prices reflect many factors other than the fair value estimates. Dietrich et al. (2001) subsequently use a direct approach to investigate the reliability of mandatory annual fair value appraisal estimates by chartered surveyors for UK investment properties and find that appraisal estimates understate actual selling prices but are considerably less biased and more accurate measures of selling price than respective historical costs. Dietrich et al. (2001) also find that the reliability of appraisal estimates increases when monitored by external appraisers and Big Six auditors.

The New Zealand (hereafter NZ) SSAP No. 17 'Accounting for Investment Properties and Properties Intended for Sale' (NZSA, 1989) previously allowed NZ companies the choice of recognising unrealised gains or losses either in the income statement, or as movements in an investment property revaluation reserve, unless the total of the reserve was insufficient to cover a deficit, in which case the amount of deficit was to be charged in the income statement as part of operating results. The NZ equivalent of IAS 40 came into effect on 1 January 2005, resulting in the elimination of the choice of recognising unrealised gains in the revaluation reserve. Owusu-Ansah and Yeoh (2006) investigate the relative value-relevance of the two alternative accounting treatments for unrealised gains on investment properties, based on a sample of NZ companies over the period 1990 to 1999, when the choice was still available. Their results show that recognition of unrealised gains in the income statement is not superior to recognition of unrealised gains in the revaluation reserve in terms of their value-relevance. However, Owusu-Ansah and Yeoh (2006) include only companies with positive changes in the value of their investment properties.

Taken together, findings from prior studies of firms in the US, UK and Australian capital markets during the 1990s suggest that investors have been provided with fair value information (whether recognised or disclosed) that is generally reliable and relevant (whether fair value estimated by management or independent valuer). More research should be undertaken to test empirically whether relevance and reliability improve after the implementation of the fair value standards on financial instruments (e.g. IAS 39) and with the extension of fair value accounting to non-financial assets (i.e. IAS 40).

Like Owusu-Ansah and Yeoh (2006), this study examines the extension of fair value accounting to investment properties and the presentation of their fair value changes in the income statements (rather than in the revaluation reserve) in particular. Unlike Owusu-Ansah and Yeoh (2006), this study employs data from accounting periods when the related fair value accounting standard HKAS 40 is implemented. Comparison is then made with those from the immediate pre-implementation accounting periods when SSAP 13 (2000) was in effect. Also, unlike Owusu-Ansah and Yeoh (2006), this study includes companies with both increases and decreases in fair values and uses a return model adapted from Easton and Harris (1991), Amir et al. (1993) and the earnings capitalisation approach[6] from Barth (1994).

[6] Two approaches are developed and used in Barth (1994), the market value approach and the earnings capitalisation approach. The former examines balance sheet items, the latter examines income statement items. Owusu-Ansah and Yeoh (2006) use a modified version of the market value approach, which is however not considered in this study. The market value approach is based on a valuation model where the market value of equity equals a firm's assets minus its liabilities on its balance sheet (Barth, 1994). The market value approach is useful if the asset being studied is reported differently in different balance sheets, for example, at historical cost or fair value because of the choice permitted for investment securities in Barth (1994). In this study, investment properties are reported in the balance sheet either under SSAP 13 (2000) at open market values or under HKAS 40 (2004) at fair values. Similar amounts will be reported in the balance sheets under SSAP 13 (2000) and HKAS 40 (2004). On the other hand, the earnings figures in the income statements are different, depending on which accounting standard is followed, which explains why this study has chosen to use the earnings capitalisation model.

## 3. Research method

Data for this study are collected from two sources. The financial statement data and results announcement dates are derived from the published financial statements downloaded from http://www.hkex.com.hk. The Morgan Stanley Capital International (hereafter MSCI) Equity Hong Kong Index is used in this study as the proxy for equity market return in Hong Kong. The equity market data including the MSCI Equity Hong Kong Index are obtained from Reuters 3000Xtra.

Companies are eligible for selection if throughout the sample period 2003–2006: (1) they are in the property industry; (2) they are listed in the Main Board of the Hong Kong Stock Exchange; (3) they report investment properties in their financial statements; and (4) there is no change of accounting year-end. These sampling criteria yield an initial sample of 92 companies.

Each company in the sample is studied twice, first in its final accounting period using SSAP 13 (2000), and then in its first accounting period using HKAS 40 (2004).

Two approaches are used in this study, the short-window event study approach and the longer window return-earnings associations approach.

In both approaches, three[7] control variables are added to control for their potentially confounding effects, including firm size and leverage which are commonly controlled for in value-relevance studies (e.g. Fan and Wong, 2002). Market-wide property price level in Hong Kong as proxied by the Centa-City Index[8] is also included as the analysis in this study involves comparison of data collected across different time periods.

### *3.1. Short-window event study*

The short-window event study approach is based on the information or signalling perspective (Beaver, 1981) on the decision usefulness of financial reporting, a perspective that has dominated financial accounting theory and research since Ball and Brown (1968). The information perspective posits that investors, in response to the release of a company's financial statements, will analyse the statements for new and unexpected information and revise their beliefs about the company's future performance, causing movements in the company's share prices and resulting in abnormal returns to investors (Ball and Brown, 1968). Thus financial statement information is considered decision-useful if it is new and unexpected, and results in abnormal returns to the investors. Easton and Harris (1991) confine financial information to reported income or earnings from the income statement and develop a model relating earnings level and earnings change to abnormal return, as follows:

$$R_{jt} - E[R_{jt}] = a_1 + a_2 \{A_{jt} / P_{jt-1}\}$$
$$+ a_3\{(A_{jt} - A_{jt-1}) / P_{jt-1}\} + e_{jt}$$

where:

$R_{jt}$ is the return on security j at time t,

$E[R_{jt}]$ is the expected return on security j at time t,

$A_{jt}$ is the accounting earnings per share of firm j over the time period t-1 to t,

$A_{jt-1}$ is the accounting earnings per share of firm j over the prior time period t-2 to t-1, and

$P_{jt-1}$ is the beginning-of-period security price per share at t-1

To evaluate the value-relevance of the HKAS 40 (2004)'s revised presentation of gains and losses in fair value of investment properties in the income statement, earnings are partitioned into two components: (1) earnings before gains and losses in fair value of investment properties; and (2) gains and losses in fair value of investment properties, as in Barth (1994). Both components are then scaled by total market value. In this study, the gains and losses in fair value of investment properties, as scaled by total market value, are assumed to have an expectation of zero as changes in fair value generally arise from random walk processes,[9] should be transitory in nature (Barth, 1994; Chambers et al., 2006) and represent unexpected information.

An abnormal return in this study is measured by the difference between the company's return during the period and the return on the market portfolio, i.e. $R_{jt} - R_{mt}$, where $R_{mt}$ is the return based on the market portfolio at time t, also known as the market-adjusted rate of return (Amir et al., 1993). Market-wide return is removed and the abnormal return obtained thus represents company-specific returns. Brown and Warner (1980) find the market-adjusted rate of return to perform reasonably well under a wide variety of conditions when compared

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[7] As HKAS 40 (2004) encourages, but does not mandate, external independent valuation for investment properties, and SSAP 13 (2000) only requires such a valuation at least every three years, whether or not there is an independent valuation should therefore be controlled for in this study. However subsequent data analysis shows that all the companies sampled in this study engage an external independent valuation appraiser for their investment properties during all the sample year-ends regardless of whether SSAP 13 (2000) or HKAS 40 (2004) is adopted. There is therefore no such need to include a control variable for independent valuation in this study.

[8] The Centa-City Index is a monthly index based on all transaction records as registered with the Land Registry in Hong Kong, to reflect market-wide property price levels during the month in Hong Kong.

[9] Changes in fair value of investment properties may behave as a random walk with drift, resulting in a non-zero expected value. The drift may be reflecting the change in the market-wide property price levels, which is therefore controlled for in this study.

with the more conventional market model used in Ball and Brown (1968) and other studies (e.g. Ball and Kothari, 1991).

In order to evaluate the value-relevance of a company's reporting of gains and losses in fair value of investment properties in the income statement, this study assesses how investors respond to such information when it first becomes publicly available, by observing the company's share price movements during a short window of three days surrounding the company's results announcement. In Hong Kong, the Hong Kong Stock Exchange Listing Rules require its Main Board listed companies to publish preliminary results the next business day after approval by the board of directors and in any event not later than four months after the date upon which the financial period ended. The preliminary results include announcements of balance sheets and income statements together with most of the notes to the accounts which have been audited and agreed by the companies' auditors and will be published in the annual reports. The preliminary results announcement day is therefore identified as the day when the information on the reporting of gains and losses in fair value of investment properties first becomes publicly available.

Because during a three-day short window there are relatively few company-specific events other than the results announcement, a three-day short-window association between the abnormal returns and accounting information released in the results announcement suggests that the accounting information is responsible for the abnormal return and provides new decision-useful information to investors. Other information in the preliminary results announcement might also be responsible for the abnormal return, for example, a change in company strategy. However, by including only companies from the property industry, where investment properties comprise a large proportion of company assets and fair value gains and losses, a large earnings source, the effects from other information are therefore narrowed in this short-window analysis.

The regression equation (1) below is used to evaluate the differential value-relevance of the HKAS 40 (2004)'s required reporting of gains and losses in fair value of investment properties in the income statement, using a three-day short window approach as adapted from Easton and Harris (1991), Amir et al. (1993) and Barth (1994) (omitting company j and year t subscripts). A differential intercept dummy variable AFTER is used to indicate whether HKAS 40 (2004) or SSAP 13 (2000) is adopted during the accounting year.

$$\begin{aligned} AR^S = \; & \alpha_1 + \alpha_2\,AFTER + \alpha_3\,EARNB \\ & + \alpha_4\,\Delta EARNB + \alpha_5\,IPVC \\ & + \alpha_6\,EARNB{*}AFTER \\ & + \alpha_7\,\Delta EARNB{*}AFTER + \alpha_8\,IPVC{*}AFTER \\ & + \alpha_9\,FIRM\ SIZE + \alpha_{10}\,LEVERAGE \\ & + \alpha_{11}\,CCINDEX + \varepsilon \end{aligned} \tag{1}$$

where:

$AR^S$ is the three-day buy-and-hold[10] abnormal return (adjusted for dividends and share splits), centred around the preliminary results announcement day, calculated using the market-adjusted return.

AFTER is a dummy variable to indicate whether HKAS 40 (2004) or SSAP 13 (2000) is in effect during the accounting year. AFTER is set equal to one if HKAS 40 (2004) is adopted for the first time during the accounting year, and zero otherwise.

EARNB is the earnings before gains and losses. In the case of HKAS 40 (2004) this is measured as earnings before gains and losses in fair value of investment properties. In the case of SSAP 13 (2000) this is measured as the earnings before deficits in open market value of investment properties in excess of the revaluation reserve, or surpluses in open market value of investment properties in excess of any deficits previously charged to the income statement (see footnote 3). EARNB is scaled by the total market value at the first day of the fifth month after the beginning of the accounting year.

ΔEARNB is the difference between EARNB in the current year and EARNB in the prior year, scaled by the total market value at the first day of the fifth month after the beginning of the accounting year.

IPVC is the gains and losses in fair value of investment properties recognised in the income statement as required by HKAS 40 (2004), or the investment properties' open market value increases and decreases disclosed in the notes

[10] While there is disagreement in the literature regarding the method for calculating long-run abnormal returns because of the inherent right-skewed non-normal distribution problem (Lyon et al., 1999), the choice of method is not so important for the measurement of short-run abnormal returns (Fama, 1998, Jakobsen and Voetmann, 2003). Because a horizon of three to five years is referred to in the literature as long-run, both the three-day short-window and 12-month long-window abnormal returns in this study provide short-run returns. Commonly used abnormal returns in event studies are buy-and-hold abnormal returns (BHAR) (monthly abnormal returns compounded) and cumulative abnormal returns (CAR) (monthly abnormal returns summed and averaged). Lyon et al. (1999) suggests BHAR is suitable in answering the question of whether sample firms earned abnormal returns over a particular period of time, while CAR is preferred where sample firms persistently earn abnormal monthly returns. Fewer data are collected for BHAR because the monthly returns compounded are simply the annual return; BHAR is therefore used in this study.

to the accounts, scaled by the total market value at the first day of the fifth month after the beginning of the accounting year.

FIRM SIZE is the natural logarithm of the book value of the total assets at the beginning of the accounting year.

LEVERAGE is the ratio of the book value of debt to the total assets at the beginning of the accounting year.

CCINDEX is the difference between the Centa-City Index at the end and the beginning of the accounting year, divided by the beginning index amount.

The focus of this study is on the significance of the differential slope coefficient $\alpha_8$ of the interaction variable IPVC*AFTER which reflects the significance of the differential causal effect of the change in the accounting standard requirement. A positive and significant $\beta_8$ gives support for the higher informativeness of the changes introduced by HKAS 40 (2004) in the presentation of changes in fair value of investment properties.

The coefficients $\alpha_6$ and $\alpha_7$ of EARNB*AFTER and ΔEARNB*AFTER reflect the differential causal effect of the HKAS 40 revision on earnings and earnings change amounts before the open market value/fair value change of investment properties (i.e. EARNB and ΔEARNB). Because the HKAS 40 revision is not concerned with such earnings and earnings change, coefficients $\alpha_6$ and $\alpha_7$ should be of no significance. However, as a result of the full convergence of HKFRS with IFRS in Hong Kong in 2005, there are other financial reporting requirement changes taking place concurrently with HKAS 40 (2004). These other reporting changes may interact with EARNB and ΔEARNB to cause an effect on the abnormal returns. The sign and significance of a6 and a7 reflect the differential causal value relevance of these other reporting changes. They are, however, outside the scope of this study and no prediction is made about them.

The coefficients $\alpha_3$ and $\alpha_4$ represent the effects of EARNB and ΔEARNB in three-day short windows when SSAP 13 (2000) is adopted in the financial statements (i.e. when AFTER equals zero). Although most related prior studies find earnings and/or earnings change to be significant in explaining abnormal returns, they do not measure the effects over short windows. When short windows are used, the results are mixed. For example, Amir et al. (1993) finds no market reaction to either earnings or earnings change around five-day announcement windows (and interprets this as reflecting an inability to differentiate the news clearly), while Haw et al. (1999) report significant price reaction using three-day windows. Also, because EARNB and ΔEARNB in this study are earnings and earnings change before investment properties' value change, they are less likely to be transitory (i.e. less likely to be 'surprises') and may show a weaker relationship with the abnormal returns. In contrast, $\alpha_5$ for IPVC may show a stronger relationship. Results from Barth et al. (1990 and 1994) give support to these predictions.

*3.2. Long-window abnormal return – unexpected earnings association*

This study also extends the returns window to a longer period of 12 months. While a three-day short window association between abnormal return and accounting information suggests that the latter is causing the former, this is not a valid claim when the association is evaluated over a long window. A longer window opens the returns up to a host of other value-relevant events. Share prices reflect all available information, not just the accounting information. As a result, it cannot be claimed that the reported accounting information causes the abnormal return during the 12-month-long window period. The most that can be claimed is that the reported accounting information and the abnormal return are associated.

The regression equation (2) for the long window is basically the same as equation (1) for the short window, except abnormal return is measured over 12 months rather than three days.

$$\begin{aligned} AR^L = \; & \beta_1 + \beta_2\,\text{AFTER} + \beta_3\,\text{EARNB} \\ & + \beta_4\,\Delta\text{EARNB} + \beta_5\,\text{IPVC} \\ & + \beta_6\,\text{EARNB*AFTER} \\ & + \beta_7\,\Delta\text{EARNB*AFTER} + \beta_8\,\text{IPVC*AFTER} \\ & + \beta_9\,\text{FIRM SIZE} + \beta_{10}\,\text{LEVERAGE} \\ & + \beta_{11}\,\text{CCINDEX} + \varepsilon \end{aligned} \tag{2}$$

where:

$AR^L$ is the buy-and-hold[11] abnormal return (adjusted for dividends and share splits), for the 12 months beginning the first day of the fifth month after the beginning of the accounting year, calculated using market adjusted return.

The Hong Kong Stock Exchange Listing Rules allow the Hong Kong Main Board listers to publish their annual financial statements (as part of the preliminary results announcements) at the latest by the last day of the fourth month after the accounting year-end. The 12-month window in this study therefore corresponds to the 12 months ending on this day (i.e. 12 months beginning eight months prior to and ending four months after the accounting year).

As with the short window regression, the higher value-relevance in the longer run of the HKAS 40

[11] See footnote 10.

**Table 1**
**First time adoption of HKAS 40 (2004)**

| *Financial statements year-end* | *No. of companies adopting HKAS 40 (2004) for the first time* |
|---|---|
| 31 December 2004 | 3 [e] |
| 31 March 2005 | 4 [e] |
| 30 April 2005 | 1 [e] |
| 30 June 2005 | 7 [e] |
| 31 July 2005 | 2 [e] |
| 31 December 2005 | 45 |
| 31 March 2006 | 22 |
| 30 April 2006 | 1 |
| 30 June 2006 | 5 |
| 31 July 2006 | 2 |
| **Total** | **92** |

[e] Early adoption of HKAS 40 (2004)

(2004) presentation of gains and losses in fair value of investment properties in the income statement is assessed by a positive and significant value of $\beta_8$.

Predictions similar to those for the earlier three-day short window (or no prediction) are made for the other variables, except stronger relationships are expected because association rather than causation is studied in 12-month windows.

## 4. Empirical results

Although HKAS 40 (2004) allows a free choice between cost and fair value models, all 92 companies in the initial sample chose to adopt the fair value model.

All the 92 companies are retained for data analysis, with extreme variable values verified against their sources. Since no procedural errors or extraordinary events are identified, all the data collected for the 92 companies are retained for the subsequent analysis.[12]

Each company is evaluated twice, in two consecutive accounting years before and after the adoption of HKAS 40 (2004).

Table 1 describes the distribution of accounting year-ends, years of last-time following of SSAP 13 (2000) and years of first-time adoption of HKAS 40 (2004) for the 92 companies in this study. Appendix A details their identities. Most companies have March 31 or December 31 accounting year-ends, and adopt HKAS 40 (2004) for the first time in 2005 or 2006. While HKAS 40 (2004) mandates adoption for annual periods beginning on or after 1 January 2005, 17 companies choose to adopt HKAS 40 (2004) early.[13]

Tables 2A and 2B contain descriptive statistics for the 92 sample companies in the study during the year(s) when HKAS 40 (2004) is adopted for the first time compared to the year(s) when SSAP 13 (2000) is adopted for the last time.

On the whole, when companies apply HKAS 40 (2004) for the first time, they are experiencing higher earnings and higher market values and offering their investors higher abnormal returns; this may be attributable to the strong economy in Hong Kong in 2005 and 2006. The Centa-City Index has indeed been increasing during the sample period, although at a significantly lower rate when HKAS 40 (2004) is applied for the first time. Firm size and Centa-City index changes are both controlled for in this study. Also all independent variables in this study are scaled by the company's beginning market value. Results show a significant increase in the proportion of investment properties relative to total assets, from 0.345 when SSAP 13(2000) is applied to 0.403 when HKAS 40 (2004) is applied.

There is also an indication of higher earnings volatility[14] as a result of applying HKAS 40 (2004). The mean gains and losses in fair value of investment properties are HK$827m which is almost equal to the earnings before such gains and losses of HK$848m. In contrast, the mean investment properties open market value excess deficits or surpluses of HK$24m amounts to only 3% of

[12] Unreported findings show that similar results are obtained if all extreme variable values are excluded.

[13] Unreported t-test results show early adopters to have significantly higher mean amounts of total assets, investment properties, total earnings and market values. Unreported regression results show similar results when a dummy variable to indicate early adopters is included in the regression equations.

[14] The increased volatility may be limited to our sample period (2004–2006). As volatility should be a function of year-over-year change, we need to examine more years in order to conclude whether or not an increase in volatility has resulted from applying HKAS 40 (2004).

**Table 2A**
**Descriptive statistics for financial statement and market data (2004–2006)**
**92 companies**

| | N | Mean | SD | Max | Min |
|---|---|---|---|---|---|
| **Selected market and economic variables** | | | | | |
| *When HKAS 40 (2004) is adopted for the first time* | | | | | |
| Company's market value (HK$m) | 92 | 9,511 | 25,642 | 178,887 | 41 |
| Centa-City Index change | 92 | 0.067 | 0.081 | 0.315 | –0.038 |
| *When SSAP 13(2000) is adopted for the last time* | | | | | |
| Company's market value (HK$m) | 92 | 7,488 | 22,623 | 174,070 | 21 |
| Centa-City Index change | 92 | 0.278 | 0.096 | 0.459 | 0.072 |
| *Paired sample t-tests of mean differences (HKAS 40 (2004) over SSAP 13) (t-value in parenthesis)* | | | | | |
| Company's market value (HK$m) | 92 | 2,023 *** | | | |
| | | (4.114) | | | |
| Centa-City Index change | 92 | –0.211 *** | | | |
| | | (–20.723) | | | |
| **Selected financial statement variables** | | | | | |
| *When HKAS 40 (2004) is adopted for the first time* | | | | | |
| Total assets (HK$m) | 92 | 15,996 | 34,459 | 196,720 | 95 |
| Investment properties (HK$m) | 92 | 6,069 | 15,068 | 116,733 | 0 |
| Investment properties to total assets | 92 | 0.403 | 0.348 | 1.547 | 0 |
| Debt to total assets | 92 | 0.182 | 0.149 | 0.790 | 0 |
| Total earnings | 92 | 1,675 | 3,478 | 20,038 | –571 |
| *When SSAP 13(2000) is adopted for the last time* | | | | | |
| Total assets (HK$m) | 92 | 14,837 | 32,684 | 198,860 | 43 |
| Investment properties (HK$m) | 92 | 5,022 | 12,654 | 100,775 | 0 |
| Investment properties to total assets | 92 | 0.345 | 0.308 | 1.289 | 0 |
| Debt to total assets | 92 | 0.162 | 0.156 | 0.536 | 0 |
| Total earnings | 92 | 1,055 | 2,569 | 18,180 | –213 |
| *Paired sample t-tests of mean differences (HKAS 40 (2004) over SSAP 13) (t-value in parenthesis)* | | | | | |
| Total assets (HK$m) | 92 | 1,159 *** | | | |
| | | (3.662) | | | |
| Investment properties (HK$m) | 92 | 1,047 *** | | | |
| | | (3.704) | | | |
| Investment properties to total assets | 92 | 0.058 ** | | | |
| | | (2.405) | | | |
| Debt to total assets | 92 | 0.020 | | | |
| | | (1.583) | | | |
| Total earnings | 92 | 620 *** | | | |
| | | (4.049) | | | |

[a]IPFVGL: Gains and losses in fair value of investment properties (reported in income statement under HKAS 40 (2004))
[b]IPOMVEDS: Investment properties' open market value excess deficits and surpluses (reported in income statement under SSAP 13)
[c]IPOMVC: Investment properties' open market value changes (disclosed in notes to the accounts under SSAP 13)
* Significant at the 10% level
** Significant at the 5% level
*** Significant at the 1% level

**Table 2B**
**Descriptive statistics for regression variables (2004–2006)**
**92 companies**

| | N | Mean | SD | Max | Min |
|---|---|---|---|---|---|
| **Regression variables** | | | | | |
| *When HKAS 40 (2004) is adopted for the first time* | | | | | |
| Short-window abnormal return | 92 | –0.005 | 0.063 | 0.217 | –0.191 |
| Long-window abnormal return | 92 | 0.156 | 0.543 | 1.911 | –0.997 |
| Earnings before IPFVGL[a] (HK$m) | 92 | 848 | 2,125 | 11,794 | –1,897 |
| Earnings change before IPFVGL[a] (HK$m) | 92 | 105 | 755 | 3,184 | –2,851 |
| IPFVGL[a] (HK$m) | 92 | 827 | 2,052 | 12,982 | –20 |
| *When SSAP 13 (2000) is adopted for the last time* | | | | | |
| Short-window abnormal return | 92 | –0.013 | 0.064 | 0.175 | –0.372 |
| Long-window abnormal return | 92 | 0.152 | 0.563 | 2.649 | –1.158 |
| Earnings before IPOMVEDS[b] (HK$m) | 92 | 757 | 1,773 | 10,601 | –216 |
| Earnings change before IPOMVEDS[b] (HK$m) | 92 | 316 | 847 | 6,777 | –459 |
| IPOMVEDS[b] (HK$m) | 92 | 24 | 146 | 886 | –784 |
| IPOMVC[c] (HK$m) | 92 | 312 | 1,110 | 8,139 | –784 |
| *T-tests of mean differences (HKAS 40 (2004) over SSAP 13) (t-value in parenthesis)* | | | | | |
| Short-window abnormal return | 92 | 0.008 (0.94) | | | |
| Long-window abnormal return | 92 | 0.004 (0.06) | | | |
| Earnings before IPFVGL[a] / IPOMVEDS[b] (HK$m) | 92 | 91 (1.02) | | | |
| Earning change before IPFVGL[a] / IPOMVEDS[b] (HK$m) | 92 | –211 * (–1.66) | | | |
| IPFVGL[a] / IPOMVC[c] (HK$m) | 92 | 515 *** (3.793) | | | |

[a]IPFVGL: Gains and losses in fair value of investment properties (reported in income statement under HKAS 40 (2004))
[b]IPOMVEDS: Investment properties' open market value excess deficits and surpluses (reported in income statement under SSAP 13)
[c]IPOMVC: Investment properties' open market value changes (disclosed in notes to the accounts under SSAP 13)
* Significant at the 10% level
*** Significant at the 1% level

**Table 3**
**Earnings volatility**

| | N | Mean | SD | Max | Min |
|---|---|---|---|---|---|
| *Applying HKAS 40 (2004)* | | | | | |
| Earnings volatility (number of standard deviations) | 92 | 0.947 | 0.728 | 1.788 | –0.167 |
| *Applying SSAP 13 (2000)* | | | | | |
| Earnings volatility (number of standard deviations) | 92 | 0.419 | 0.589 | 1.747 | –0.190 |

Paired-sample $t$-test results:

Paired differences: 0.528 number of standard deviations
$t\,(91) = 4.679$
$p = 0.000$

the earnings before such excess deficits or surpluses of HK$757m.

Further indication of higher earnings volatility[15] is available in Table 3 showing the results of a paired-sample t-test performed to compare the earnings volatility before and after the application of HKAS 40 (2004). Earnings volatility is expressed as the number of standard deviations from a five-year mean (mean of the earnings of the five years ending on the year of HKAS 40 (2004) application). Earnings volatility is significantly higher after the adoption of HKAS 40 (2004) ($t = 4.678, p = 0.000$).

*4.1. Short window event study*

Table 4 reports the regression results from the estimation of equation (1). Results for the short-window event study provide evidence that the presentation of changes in fair value of investment properties in the income statements as required by HKAS 40 (2004) is more informative to investors than the presentation required by SSAP 13 (2000). Investors respond to the information on changes in fair value in the income statement, as released in the results announcement, causing abnormal returns. The coefficient $\alpha_8$ of the interaction variable IPVC*AFTER in equation (1) is positive and significant at the 5% level ($p = 0.022$).

As expected, neither earnings (EARNB) nor earnings change before investment properties' open market value/changes in fair value (ΔEARNB) is significant in explaining the abnormal return within the short window when SSAP 13 (2000) is adopted in the financial statements.

Although the overall $R^2$ is only 1.0%, this is consistent with the results from prior short-window studies.

All the coefficients are positive except that of investment properties' value changes (i.e. IPVC), which is negative (but not statistically significant). Barth et al. (1990) and Barth (1994) also find similar negative coefficients for securities market price gains and losses and interpret them as evidence of a market that perceives that securities gains and losses are used to smooth earnings.

*4.2. Long-window abnormal return – unexpected earnings association*

The regression results for the long window abnormal return and unexpected earnings association are reported in Table 4. As the window opens wider, the earnings before changes in open market value or fair value (i.e. EARNB) in equation (2) also become significant at the 5% level and the overall adjusted $R^2$ increases to 17.7%. This is consistent with the results from prior studies using long windows where significance of earnings is found together with higher overall $R^2$.

The coefficient $\beta_8$ of the interaction variable IPVC*AFTER in equation (2) is positive and significant at the 10% level ($p = 0.069$), consistent with the result from the short window regression in this study. This result conflicts with that from Owusu-Ansah and Yeoh (2006), who find that the recognition of unrealised gains in the income statement is not superior to that in the revaluation reserve in terms of value-relevance. However, the results from this present study are more persuasive on a number of grounds: (1) Owusu-Ansah and Yeoh (2006) is based on data from the 1990s before the fair value accounting requirement came into effect, while our study is based on the actual reactions from investors to the implementation of the fair value model (i.e. HKAS 40 (2004)) compared to their previous reaction under SSAP 13 (2000); (2) this study has a cleaner test setting as we use only the earnings capitalisation approach which focuses on income statement items, whereas Owusu-Ansah and Yeoh (2006) adopt a market value approach which examines balance sheet items. This latter approach is not appropriate since the values of investment properties (open market value or fair value) reported in balance sheets are similar, before and after the change in accounting requirements (see footnote 6); (3) the sample companies in this present study belong to a single industry – property; as detailed earlier, the effects from other information are narrower if the sample is restricted to those companies drawn only from the property industry. Property companies are expected to have more precise value estimates for their investment properties, but the Owusu-Ansah and Yeoh (2006) sample includes companies from different industries; (4) the findings of this present study have greater generality, in that they embrace companies with both value increases and decreases, while Owusu-Ansah and Yeoh (2006) confine their sample to those companies with positive changes in values only.

The coefficient $\beta_6$ of the interaction variable EARNB*AFTER in equation (2) is also positive and significant at the 10% level ($p = 0.092$). This gives support for the value-relevance of other financial reporting changes that are taking place concurrently with HKAS 40 (2004) as a result of the full convergence of HKFRS with IFRS in Hong Kong in 2005. These other reporting changes interact with the earnings before changes in the fair value of open market value or fair value of investment properties (i.e. EARNB) to result in a stronger association with abnormal returns.

As with the results from equation (1) for the short window, all the coefficients are positive except those of IPVC and ΔEARNB*AFTER, which

[15] We also test the change in volatility after controlling for the change in the market-wide property price levels (measured by the Centa-City Index). Similar results are obtained.

**Table 4**
**Regression results**

| | $AR^S$ *Equation (1)* | $AR^L$ *Equation (2)* |
|---|---|---|
| Intercept | –0.101 * | –1.561 *** |
| | (–1.72) | (–3.37) |
| AFTER | 0.009 | 0.121 |
| | (0.57) | (0.99) |
| EARNB | 0.010 | 0.327 ** |
| | (0.51) | (2.19) |
| Δ EARNB | 0.000 | 0.168 |
| | (0.00) | (1.29) |
| IPVC | –0.024 | –0.076 |
| | (–1.49) | (–0.60) |
| EARNB *AFTER | 0.025 | 0.439 * |
| | (0.77) | (1.70) |
| Δ EARNB *AFTER | 0.001 | –0.230 |
| | (0.05) | (–1.34) |
| IPVC *AFTER | 0.062 ** | 0.392 * |
| | (2.31) | (1.83) |
| FIRM SIZE | 0.003 | 0.066 *** |
| | (1.28) | (3.12) |
| LEVERAGE | 0.005 | 0.124 |
| | (0.16) | (0.49) |
| CCINDEX | 0.047 | 0.693 |
| | (0.85) | (1.58) |
| *Adj.* $R^2$ | 0.010 | 0.177 |
| *N* | 184 | 184 |

$AR^S$ = three-day buy-and-hold abnormal return (adjusted for dividends and share splits), centred around the preliminary results announcement day, calculated using the market-adjusted return

$AR^L$ = buy-and-hold abnormal return (adjusted for dividends and share splits), for the 12 months beginning the first day of the fifth month after the beginning of the accounting year, calculated using the market-adjusted return

AFTER = a dummy variable to indicate whether HKAS 40 (2004) or SSAP 13 is in effect during the accounting year. AFTER is set equal to one if the new HKAS 40 (2004) is adopted for the first time during the accounting year, and zero otherwise

EARNB = earnings before gains and losses in fair value of investment properties recognised under the new HKAS 40 (2004), OR the earnings before investment properties' open market deficits in excess of the revaluation reserve balance or investment properties' subsequent open market value surpluses in excess of the deficits previously charged to the income statement, recognised under the old superseded SSAP 13. EARNB is scaled by the total market value at the first day of the fifth month after the beginning of the accounting year

ΔEARNB = difference between EARNB in the current year and EARNB in the prior year, scaled by the total market value at the first day of the fifth month after the beginning of the accounting year

IPVC = gains and losses in fair value of investment properties recognised in the income statement as required by the new HKAS 40 (2004), OR the investment properties' open market value increases and decreases disclosed in the notes to the accounts, scaled by the total market value at the first day of the fifth month after the beginning of the accounting year

FIRM SIZE = natural logarithm of the book value of the total assets at the beginning of the accounting year

LEVERAGE = ratio of the book value of debt to the total assets at the beginning of the accounting year

CCINDEX = difference between the Centa-City Index at the end and at the beginning of the accounting year, divided by the beginning index amount

* significant at the 10% level
** significant at the 5% level
*** significant at the 1% level

are negative (but not statistically significant).

*4.3. Sensitivity analysis*

If the earnings partitioning approach developed in Barth (1994) is to be followed exactly in this study, the variable IPVC (changes in fair value of investment properties) should be equal to the excess open market value deficits or surpluses recognised in the income statement for the accounting year when SSAP 13 (2000) is followed. This study has chosen to use the open market value increases or decreases (as disclosed in the notes to the accounts), in order to be consistent with the fair value gains and losses that are used to measure the variable IPVC when HKAS 40 (2004) is adopted. A sensitivity analysis is conducted using the excess open market value deficits or surpluses when SSAP 13 (2000) is followed and similar results are obtained.

Another sensitivity analysis is conducted using the Hong Kong Hang Seng Composite Index as the proxy for equity market return in Hong Kong, and again, similar results are found. The Hong Kong Hang Seng Composite Index covers 90% of the market capitalisation of the shares listed on the Main Board of SEHK and there are currently 200 constituent shares in this index.

## 5. Conclusions

Part of the debate about the adoption of fair value accounting for investment properties is on the value-relevance of presenting changes in fair value in the income statement, compared to reporting such changes in the revaluation reserve. This study informs this debate by providing evidence on the value-relevance of the presentation of changes in fair value in the income statement for Hong Kong listed companies in the properties industry. Results show a significant market price reaction to investment properties' fair value change information as included in companies' annual results announcements. Results also show a significant association between the market-adjusted annual share returns and the presentation of the investment properties' fair value change in the income statement. These results strongly suggest that investors appear to place more value on HKAS 40 (2004)'s presentation of changes in fair value of investment properties in the income statement, when compared with the presentation in the revaluation reserve under SSAP 13 (2000). The results also support the existing literature on the value-relevance of presentation locations of accounting amounts in general.

The results from this study also have implications for companies around the world that prepare their financial statements using International Financial Reporting Standards.

However, because all the companies in our sample choose to adopt the fair value model, we do not have a control group of companies that do not adopt the fair value model. We cannot therefore eliminate the possibility that our results are driven by other events happening at the same time as the adoption of HKAS 40 (2004).

**Appendix A**
**92 companies in the sample (184 firm-standard)**

| | *SEHK Code* | | *Accounting year-end* | *Last time following SSAP 13* | *First time adoption of HKAS 40 (2004)* |
|---|---|---|---|---|---|
| 1 | 1 | Cheung Kong (Holdings) Ltd | 31 December | 2003 | 2004 |
| 2 | 154 | Beijing Development (Hong Kong) Ltd | 31 December | 2003 | 2004 |
| 3 | 758 | Junefield Department Store Group Ltd | 31 December | 2003 | 2004 |
| 4 | 35 | Far East Consortium International Ltd | 31 March | 2004 | 2005 |
| 5 | 172 | Goldbond Group Holdings Ltd | 31 March | 2004 | 2005 |
| 6 | 277 | Tern Properties Co Ltd | 31 March | 2004 | 2005 |
| 7 | 412 | Heritage International Holdings Ltd | 31 March | 2004 | 2005 |
| 8 | 735 | Oriental Investment Corporation Ltd | 30 April | 2004 | 2005 |
| 9 | 10 | Hung Lung Group Ltd | 30 June | 2004 | 2005 |
| 10 | 12 | Henderson Land Development Company Ltd | 30 June | 2004 | 2005 |
| 11 | 83 | Sino Land Company Ltd | 30 June | 2004 | 2005 |
| 12 | 97 | Henderson Investment Ltd | 30 June | 2004 | 2005 |
| 13 | 101 | Hang Lung Properties Ltd | 30 June | 2004 | 2005 |
| 14 | 131 | Cheuk Nang (Holdings) Ltd | 30 June | 2004 | 2005 |
| 15 | 247 | Tsim Sha Tsui Properties Ltd | 30 June | 2004 | 2005 |
| 16 | 488 | Lai Sun Development Co Ltd | 31 July | 2004 | 2005 |
| 17 | 1125 | Lai Fung Holdings Ltd | 31 July | 2004 | 2005 |
| 18 | 14 | Hysan Development Company Ltd | 31 December | 2004 | 2005 |

**Appendix A**
**92 companies in the sample (184 firm-standard)** *(continued)*

| | *SEHK Code* | | *Accounting year-end* | *Last time following SSAP 13* | *First time adoption of HKAS 40 (2004)* |
|---|---|---|---|---|---|
| 19 | 24 | Burwill Holdings Ltd | 31 December | 2004 | 2005 |
| 20 | 28 | Tian An China Investment Company Ltd | 31 December | 2004 | 2005 |
| 21 | 34 | Kowloon Development Company Ltd | 31 December | 2004 | 2005 |
| 22 | 41 | Great Eagle Holdings Ltd | 31 December | 2004 | 2005 |
| 23 | 56 | Allied Properties (HK) Ltd | 31 December | 2004 | 2005 |
| 24 | 66 | MTR Corporation Ltd | 31 December | 2004 | 2005 |
| 25 | 68 | Lee Hing Development Ltd | 31 December | 2004 | 2005 |
| 26 | 75 | Y.T. Reality Group Ltd | 31 December | 2004 | 2005 |
| 27 | 89 | Tai Sang Land Development Ltd | 31 December | 2004 | 2005 |
| 28 | 106 | Shenzhen Hing-Tech Holdings Ltd | 31 December | 2004 | 2005 |
| 29 | 115 | Grand Field Group Holdings Ltd | 31 December | 2004 | 2005 |
| 30 | 123 | Guangzhou Investment Company Ltd | 31 December | 2004 | 2005 |
| 31 | 127 | Chinese Estate Holdings Ltd | 31 December | 2004 | 2005 |
| 32 | 132 | China Investment Holdings Ltd | 31 December | 2004 | 2005 |
| 33 | 141 | Great China Holdings Ltd | 31 December | 2004 | 2005 |
| 34 | 156 | Lippo China Resources Ltd | 31 December | 2004 | 2005 |
| 35 | 169 | China Fair Land Holdings Ltd | 31 December | 2004 | 2005 |
| 36 | 171 | Silver Grant International Industries Ltd | 31 December | 2004 | 2005 |
| 37 | 173 | K. Wah International Holdings Ltd | 31 December | 2004 | 2005 |
| 38 | 184 | Keck Seng Investments (Hong Kong) Ltd | 31 December | 2004 | 2005 |
| 39 | 201 | Magnificent Estates Ltd | 31 December | 2004 | 2005 |
| 40 | 219 | Shun Ho Technology Holdings Ltd | 31 December | 2004 | 2005 |
| 41 | 230 | Onfem Holdings Ltd | 31 December | 2004 | 2005 |
| 42 | 242 | Shun Tak Holdings Ltd | 31 December | 2004 | 2005 |
| 43 | 251 | SEA Holdings Ltd | 31 December | 2004 | 2005 |
| 44 | 257 | China Everbright International Ltd | 31 December | 2004 | 2005 |
| 45 | 258 | Tomson Group Ltd | 31 December | 2004 | 2005 |
| 46 | 286 | G-Prop (Holdings) Ltd | 31 December | 2004 | 2005 |
| 47 | 366 | Luks Industrial (Group) Ltd | 31 December | 2004 | 2005 |
| 48 | 373 | Allied Group Ltd | 31 December | 2004 | 2005 |
| 49 | 431 | Greater China Holdings Ltd | 31 December | 2004 | 2005 |
| 50 | 588 | Beijing North Star Co Ltd | 31 December | 2004 | 2005 |
| 51 | 604 | Shenzhen Investment Ltd | 31 December | 2004 | 2005 |
| 52 | 617 | Paliburg Holdings Ltd | 31 December | 2004 | 2005 |
| 53 | 635 | Playmates Holdings Ltd | 31 December | 2004 | 2005 |
| 54 | 649 | Shimao International Holdings Ltd | 31 December | 2004 | 2005 |
| 55 | 683 | Kerry Properties Ltd | 31 December | 2004 | 2005 |
| 56 | 688 | China Overseas Land & Investment Ltd | 31 December | 2004 | 2005 |
| 57 | 730 | Shougang Concord Grand (Group) Ltd | 31 December | 2004 | 2005 |
| 58 | 878 | Soundwill Holdings Ltd | 31 December | 2004 | 2005 |
| 59 | 898 | Multifield International Holdings Ltd | 31 December | 2004 | 2005 |
| 60 | 1200 | Midland Holdings Ltd | 31 December | 2004 | 2005 |
| 61 | 1207 | Shanghai Real Estate Ltd | 31 December | 2004 | 2005 |
| 62 | 2355 | Baoye Group Company Ltd | 31 December | 2004 | 2005 |
| 63 | 20 | Wheelock and Company Ltd | 31 March | 2005 | 2006 |
| 64 | 22 | Mexan Ltd | 31 March | 2005 | 2006 |
| 65 | 49 | Wheelock Properties Ltd | 31 March | 2005 | 2006 |
| 66 | 88 | Tai Cheung Holdings Ltd | 31 March | 2005 | 2006 |
| 67 | 129 | Asia Standard International Group Ltd | 31 March | 2005 | 2006 |
| 68 | 160 | Hon Kwok Land Investment Company Ltd | 31 March | 2005 | 2006 |
| 69 | 163 | Emperor International Holdings Ltd | 31 March | 2005 | 2006 |
| 70 | 164 | Premium Land Ltd | 31 March | 2005 | 2006 |
| 71 | 166 | New Times Group Holdings Ltd | 31 March | 2005 | 2006 |

**Appendix A**
**92 companies in the sample (184 firm-standard)** *(continued)*

| | *SEHK Code* | | *Accounting year-end* | *Last time following SSAP 13* | *First time adoption of HKAS 40 (2004)* |
|---|---|---|---|---|---|
| 72 | 202 | Interchina Holdings Co Ltd | 31 March | 2005 | 2006 |
| 73 | 213 | National Electronics Holdings Ltd | 31 March | 2005 | 2006 |
| 74 | 224 | Pioneer Global Group Ltd | 31 March | 2005 | 2006 |
| 75 | 237 | Safety Godown Co Ltd | 31 March | 2005 | 2006 |
| 76 | 278 | Wah Ha Realty Company Ltd | 31 March | 2005 | 2006 |
| 77 | 287 | Winfair Investment Company Ltd | 31 March | 2005 | 2006 |
| 78 | 298 | Chuang's China Investments Ltd | 31 March | 2005 | 2006 |
| 79 | 480 | HKR International Ltd | 31 March | 2005 | 2006 |
| 80 | 499 | Hycomm Wireless Ltd | 31 March | 2005 | 2006 |
| 81 | 711 | Chun Wo Holdings Ltd | 31 March | 2005 | 2006 |
| 82 | 1036 | Winsor Properties Holdings Ltd | 31 March | 2005 | 2006 |
| 83 | 1124 | Costal Greenland Ltd | 31 March | 2005 | 2006 |
| 84 | 2340 | Synergies Holdings Ltd | 31 March | 2005 | 2006 |
| 85 | 288 | Berjaya Holding (HK) Ltd | 30 April | 2005 | 2006 |
| 86 | 16 | Sun Hung Kai Property Ltd | 30 June | 2005 | 2006 |
| 87 | 17 | New World Development Co Ltd | 30 June | 2005 | 2006 |
| 88 | 54 | Hopewell Holdings Ltd | 30 June | 2005 | 2006 |
| 89 | 659 | NWS Holdings Ltd | 30 June | 2005 | 2006 |
| 90 | 917 | New World China Land Ltd | 30 June | 2005 | 2006 |
| 91 | 193 | Capital Estate Ltd | 31 July | 2005 | 2006 |
| 92 | 1191 | China Rich Holdings Ltd | 31 July | 2005 | 2006 |

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# Earnings quality in ex-post failed firms

**Juan Manuel García Lara, Beatriz García Osma and Evi Neophytou***

**Abstract**—This paper analyses earnings quality in ex-post failed firms. Using a large sample of UK bankrupt firms, we find that failed firms manage earnings upwards in the four years prior to failure. This manipulation is achieved in two ways: (1) through accounting (accruals) manipulation; and (2) by implementing real operating actions that deviate from normal practice. We show that these two types of manipulation lead to reduced earnings reliability. We use conditional conservatism as a proxy for reliability, as prior literature links conditional accounting conservatism to better governance and positive economic outcomes. Our results show that conditional conservatism decreases substantially in the years prior to failure. Finally, we show that accruals manipulation is more pronounced in ex-post bankrupt firms with low ex-ante probability of failure, and that ex-post bankrupt firms with high ex-ante failure probability, having likely exhausted the opportunities for accrual manipulation, manipulate real operations more aggressively.

**Keywords:** firm failure, accruals management, real earnings management, conditional conservatism. earnings quality, bankruptcy

## 1. Introduction

We analyse earnings quality for a large sample of UK failed firms in the four years before failure (from year *t*-4 to year *t*, where *t* is the year of firm failure). While there is extensive research analysing the quality of accounting numbers in financially troubled firms, empirical evidence on the earnings behaviour of ex-post bankrupt (failed) firms in the years leading up to bankruptcy is scarce. Prior research has focused exclusively on accounting manipulation either in the year immediately prior to failure (Smith et al., 2001) or aggregating information from the years before failure (Rosner, 2003). These studies do not consider how manipulation evolves as failure approaches, and report only descriptive evidence of real activities manipulation. In addition, they fail to provide evidence on the consequences that managerial manipulation has over the reliability or usefulness of earnings.[1]

We argue that in the presence of deteriorating firm performance: (1) managers of ex-post failed firms manipulate both accruals and real activities to conceal poor performance in the years preceding failure; however, (2) as failure approaches, and when the opportunities for further accrual manipulation are exhausted, it is anticipated that prior manipulation will reverse into large negative accruals. Hence, prior studies in this area that aggregate data from the final years prior to bankruptcy or analyse only the year before failure, potentially provide confounding evidence on earnings quality patterns in ex-post failed firms. Generally, we expect that managers prefer the manipulation of accruals over the manipulation of real activities, as the latter is more costly. Consequently, (3) managers are likely to resort to the manipulation of real activities only when there is limited scope left for accrual manipulation. Finally, we argue that (4) the manipulation of both accruals and real activities has severe consequences on the reliability of earnings for decision-making. We expect that managerial manipulation reduces the reliability of accounting numbers, leading to reduced conditional conservatism. Therefore, we explore and provide new evidence on three important aspects that have not been analysed in detail by prior work, namely managerial use of accounting and real activities manipulation to hide poor performance, how manipulation evolves as failure looms closer and the opportunities for successfully postponing bankruptcy diminish,

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[1] To be consistent with international literature, we use the term 'bankruptcy' to refer to firm failure. However, in the UK this term is only appropriate for individuals. According to the UK *Insolvency Act 1986*, the correct term to use for companies is insolvency (see www.insolvency.gov.uk). We use 'failure', 'insolvency' and 'bankruptcy' interchangeably throughout the paper.

and the consequences of this manipulation on the reliability of earnings.

Furthering our understanding of these three issues is crucial for: (1) the adequate development of new accounting standards and corporate governance regulations aimed at tackling opportunistic managerial behaviour in financially troubled firms; as well as (2) the improvement of bankruptcy prediction models, by explicitly taking into account managerial attempts to hide poor performance. It is commonly accepted that investors can predict bankruptcy fairly accurately using prediction models. However, these models are based to a great extent on accounting data that can be manipulated by management (Ohlson, 1980). We show that managers make both accounting and operating decisions that affect the reliability of accounting numbers, and how these decisions accumulate and reverse in the year prior to bankruptcy. This evidence is potentially important in designing bankruptcy prediction models that may be able to account for the observations reported in this paper. Finally, we provide limited evidence (3) that managers manipulate real decisions to postpone bankruptcy. This evidence raises awareness of sub-optimal decision making that may be particularly relevant for debt holders and claimants. Sub-optimal operating, financing or investment decision-making may deplete the firm of some of its assets or, at the very least, reduce their liquidation value. The evidence on real actions is of interest to those directly involved in the liquidation process of the company, in terms of timing the decision on when to initiate a winding-up petition, or to best estimate the liquidation value of the firm prior to bankruptcy.

Using a large sample of continuing and failed UK firms from 1995 to 2004, we find that ex-post failed firms engage in both accounting manipulation and real activities manipulation in the four years prior to failure. This is consistent with the evidence in Beaver (1966) that the properties of accounting numbers between failing and continuing firms differ in the five years prior to the failure event. As bankruptcy approaches, the manipulation unravels and we observe large negative accruals in the year just before failure. We analyse the existence of accounting and real activities manipulation using discretionary accruals models (Jones, 1991; Dechow et al. 1995; Kasznik, 1999) and the abnormal cash flow model proposed by Dechow et al. (1998) as implemented by Roychowdhury (2006). Our results provide evidence consistent with the existence of managerial preferences and trade-offs in selecting earnings management instruments as suggested – but untested – by prior research (e.g. Peasnell et al., 2000). We show that firms with a higher probability of bankruptcy, likely having exhausted their opportunities for successful accounting manipulation, engage in the manipulation of real activities. Our sample of failed firms provides a good setting for a test of real earnings management, because even if real manipulation implies higher costs, bankruptcy is certainly more costly. Finally, our results are consistent with managers successfully concealing poor performance through accounting manipulation, as we find that accruals manipulation is more pronounced in ex-post failed firms with low ex-ante failure probability.

In what we think is a significant contribution of the paper, we formally assess the impact of these two types of manipulation on the reliability or usefulness of earnings. To do so, we analyse if ex-post failed firms report significantly less conditionally conservative earnings than continuing firms. Following Basu (1997), we identify conditional conservatism with an asymmetric recognition of economic gains and losses into earnings, and we measure it using the methodology proposed by Ball and Shivakumar (2005).[2] Prior research shows that conditional conservatism is a desirable property of earnings as it is a key corporate governance provision that provides important economic benefits such as reducing agency problems driven by information asymmetries (Beekes et al., 2004; Ahmed and Duellman, 2007; García Lara et al., 2007, 2009; LaFond and Watts, 2008), and that it varies substantially across firms in a predictable way (e.g. Ball et al., 2000; Qiang, 2007), especially when there are differences in accrual-based earnings management (García Lara et al., 2005). Our results show that ex-post bankrupt firms report less conditionally conservative earnings in the years prior to failure than continuing firms. This reduced conditional conservatism, as suggested by prior research, increases agency problems and the probability of stakeholder expropriation.

The paper makes several contributions to the existing literature. Specifically, we add to the stream of research on the quality of earnings in failed firms in several ways. First, we analyse two different aspects of earnings management: accounting manipulation and real activity manipulation, showing that managers resort to both. We also provide evidence on the timing of the manipulation,

[2] We use the term conditional conservatism following Beaver and Ryan (2005). Other studies use 'earnings conservatism', 'income-statement conservatism', 'ex-post conservatism' or 'news-driven' conservatism to refer to the same phenomenon.

[3] Prior research assumes the manipulation is the same from five years prior to bankruptcy to the year just before bankruptcy (Rosner, 2003), or analyses only the manipulation in the year just before failure (Smith et al., 2001). By aggregating data over the years prior to failure, an analysis would similarly classify a firm that steadily increases earnings quality from low to high with a firm that does the reverse; as on average across the five years they would both look similar when they are indeed very different.

as we show it starts four years prior to failure, and that accrual manipulation unravels in the year just before failure.[3] Second, we show ex-post failed firms with a low ex-ante failure probability engage in more pronounced accruals manipulation. We interpret this evidence as firms being successful in hiding poor performance through accruals manipulation. On the other hand, firms with a high ex-ante failure probability, having likely exhausted their opportunities for accrual manipulation, manipulate real activities more aggressively. Third, we conduct conditional conservatism tests to analyse whether the manipulation (both accounting and real activities manipulation) leads to more aggressive accounting. We find that ex-post failed firms report less conditionally conservative earnings and, as suggested by prior research in conditional conservatism, this implies increased agency problems. Finally, prior research has focused on the US (Rosner, 2003), or in periods of recession (Smith et al., 2001). We use a UK sample over the period 1995–2004. The UK insolvency code allows for a wider definition of bankruptcy, with different implications, than in the US (Franks et al., 1996; Bradbury, 2007).

The remainder of the paper is structured as follows. Section 2 reviews the existing literature in this area and presents the research questions. Section 3 details the sample selection procedure and describes the methodology. Section 4 discusses the empirical results and Section 5 concludes.

## 2. Background and research questions

### *2.1. Firm failure in the UK*

The UK *Insolvency Act 1986* and its subsequent amendments govern the legal failure routes, rules and regulations for insolvent British companies. The Act provides several legal courses of action for companies in financial distress, the most common and popular ones being administration, administrative receivership and liquidation.

An administration order is a court order placing a company that is, or is likely to become, insolvent under the control of an administrator following a petition by the company, its directors or a creditor (Part II, *Insolvency Act 1986*, s. 8–27). The purpose of the order is to preserve the company's business and assets to allow a reorganisation, or ensure the most advantageous realisation of its assets whilst protecting it from action by its creditors.

An administrative receivership arises as a result of a company defaulting on secured borrowing (Part III, *Insolvency Act 1986*, s. 28–72). The borrowing is usually from a bank. Under receivership, a receiver is appointed by the creditor to run the company with the objective of recovering the outstanding bank finance, through any means (Chapter I, Part III). From a company point of view, the company is rarely saved in its existing form, its assets are often subject to a meltdown and often jobs and economic activity is lost.

Finally, the liquidation or winding up of companies is the procedure whereby the assets of a company are gathered in and realised, the liabilities are met, and surplus, if any exists, is distributed to members (Part IV, *Insolvency Act 1986*, s. 73–219). Liquidation can be either compulsory or voluntary. In compulsory liquidation, a winding-up petition is initiated, usually by a creditor, due to an outstanding debt that the company has not paid (Chapter VI, Part IV, *Insolvency Act 1986*). Voluntary liquidation is the placing of the company into liquidation by resolution of its members. There are two types of voluntary liquidation: creditors' voluntary liquidation and members' voluntary liquidation. In the former case, the directors of the company, who assess that the company is, or will be, insolvent, approach an insolvency practitioner to wind up the company. In such cases, the creditors have the right to reject (by value of vote) the choice of the liquidator chosen by the company directors (Chapter IV, Part IV, *Insolvency Act 1986*). On the other hand, a members' voluntary liquidation does not involve insolvency and comes about merely because the company shareholders wish to have the value of their holding realised (Chapter III, Part IV, *Insolvency Act 1986*).

### *2.2. Prior evidence on earnings management in financially distressed firms*

Prior research on earnings management by troubled companies commonly focuses on firms that are showing signs of financial distress. A popular proxy for distress is debt covenant violation. Positive accounting theory predicts that firms approaching covenant violations will make income-increasing accounting choices to loosen their debt constraints (Watts and Zimmerman, 1986). Sweeney (1994) and DeFond and Jiambalvo (1994) find evidence consistent with this hypothesis. Nonetheless, results are not always consistent; in fact, several studies find firms in distress reporting large negative accruals (DeAngelo et al., 1994; Peltier-Rivest, 1999; Saleh and Ahmed, 2005), and argue that managers prefer to reflect the firm's financial troubles to signal their willingness to deal with them (DeAngelo et al., 1994), or to obtain concessions from labour unions and subsidies from government. In all likelihood, a number of these financially troubled firms will approach or enter into technical default. However, covenant violations are not always associated with increased distress risk. To account for this, several studies employ alternative definitions of distress such as dividend declines and series of consecutive losses (DeAngelo et al., 1994; Peltier-Rivest, 1999), or receiving going-concern audit opinions (Butler et al., 2004). These studies find that distressed firms

have large negative, performance-related accruals, but do not attribute this finding to earnings management practices. Instead, they argue that firms with high risk of failure engage in liquidity enhancing transactions in an attempt to survive.

The proxies used by prior research are likely to overstate the actual cases of severely troubled firms. Clearly, a more objective definition of financial distress is legal (involuntary) failure. Few studies, however, make use of bankruptcy data, probably due to the inherent difficulty in finding adequate data for firms that do not exist anymore. Also, bankruptcy presents a rather narrow definition of failure, with only a very small percentage of listed firms going bankrupt. However, the stakeholders of a failed company suffer more severe losses than those of firms exhibiting weak performance. Hence, the detection of aggressive accounting practices in the years that precede bankruptcy becomes critical.

Only the prior work of Smith et al. (2001), Rosner (2003) and Charitou et al. (2007a, 2007b) directly examines the properties of earnings in actually failed firms. For an Australian sample, Smith et al. find that ex-post bankrupt firms do not engage in accounting changes to inflate earnings in the year just before failure (year *t*-1). They find that distressed firms that do not end up being bankrupt change their accounting policies to increase earnings. However, they do not directly examine whether ex-post bankrupt firms changed their accounting policies in the years before *t*-1, while managers still believe in the recovery of the firm.[4] Also, they do not separately analyse accrual accounting behaviour in bankrupt firms with a low ex-ante probability of bankruptcy. Because their benchmark for failure is set on a specific date (12 months after filing the 1988 accounts), they cannot follow up on the behaviour of their distressed company sample. In contrast, our procedure allows us to follow up on firms that show little outward signs of distress but end up bankrupt. Since we do not know if the distressed firms in the Smith et al. study ended up bankrupt, their results on accounting changes in a period of economic downturn are difficult to interpret. Additionally, these changes in policies have to be reported in the financial statements, and thus, as earnings management instruments, they lack the element of concealment.

More closely related to our study, Rosner (2003) investigates accrual manipulation on a sample of 293 failed US companies and finds that firms manipulate earnings upwards in pre-bankruptcy non-going-concern years. However, in her tests she pools all years together and assumes that the manipulation of earnings in ex-post bankrupt firms is the same during the five years prior to failure. Our results indicate that this is not the case. In addition, she shows that in the year preceding failure, firms receiving going-concern opinions exhibit income-decreasing behaviour. She attributes this to the overstatement reversals that auditors demand when they eventually detect a going-concern problem.

Finally, the work of Charitou et al. (2007a, 2007b), which closely relates to that of Rosner (2003), uses a sample of US firms that filed for Chapter 11 to analyse earnings management practices in distressed firms. However, their focus is on analysing the role played by auditors and other monitoring bodies in constraining earnings manipulation in the year just before bankruptcy and on the links between earnings manipulation in the year prior to filing for bankruptcy and the likelihood of survival thereafter.

### 2.3. *Research questions*

#### 2.3.1. *Accounting manipulation*

Firms approaching bankruptcy probably engage in income-increasing earnings management practices (decreasing cost of goods sold, releasing bad debt accruals, etc.) in an attempt to ride out what is probably deemed by management as a temporary bad period. Managerial optimism most likely prevails, leading managers to maintain the belief that firm performance will improve in the following periods and thus, to keep manipulating earnings upwards even when they know that their actions will accumulate and reverse in the future (Graham et al., 2005). However, there are limits to managerial optimistic biases. The balance sheet accumulates previous decisions, placing constraints to the number of consecutive years when over-optimism can be exerted without committing GAAP violations (Barton and Simko, 2002). Precisely because firms end up failed, there is an assurance that deteriorating financial health was not a temporary concern. It is expected that: (1) the scope for accrual management will decrease as bankruptcy approaches, as managers exhaust their opportunities for successful manipulation; and also that (2) the actions undertaken in the years leading up to bankruptcy will accumulate in the final year, cascading into large negative accruals when previous manipulation reverses and deteriorating performance cannot be hidden any longer. Thus, we test the following hypotheses (in alternative form):

**H1A:** Failed firms manage earnings in the years before failure;

**H1B:** In the year prior to bankruptcy, as managers exhaust their opportunities for

[4] Their sample covers only 1988. As they recognise, their evidence might not be generalisable due to the stock market crash of 1987 and the subsequent financial crisis (likely leading to increased monitoring). We cover a longer and more stable time period (1998–2004), in a stable macroeconomic environment (the UK).

successful accrual manipulation, previously manipulated accruals reverse, leading to an accumulation of negative accruals just before failure.

*2.3.2. Real earnings management*

Following Roychowdhury (2006), we define real earnings manipulation as 'management actions that deviate from normal business practices, undertaken with the primary objective of meeting certain earnings thresholds' (p. 336). Earnings incentives give rise to sub-optimal operating, investment and financing decision making, when managers resort to real actions to meet their income targets. The seminal work of Schipper (1989: 92) puts forward as an example of real earnings management the timing of investment or financing decisions to alter reported earnings. Subsequent research demonstrates that managers time the sale of long-lived assets and investments (Bartov, 1993), delay and abandon research and development projects, give more lenient credit terms or reduce necessary expenses such as those on advertising and asset maintenance to meet their earnings targets (Bushee, 1998; Graham et al., 2005, Roychowdhury, 2006; Cohen et al., 2008; García Osma and Young, 2009).[5] Thus, whilst real earnings management is not an entirely accounting phenomenon (in that it is not directly achieved through accounting), it is driven by the desire to keep up accounting appearances. A survey by Graham et al. (2005) of 401 financial executives indicates a preference for real actions over purely accounting decisions to manipulate earnings. Similar UK survey studies present comparable views (Demirag, 1995; Grinyer et al., 1998).

In the case of failing firms, real earnings manipulation is aimed at increasing current earnings to conceal poor performance. However, real earnings manipulation is expected to reduce future cash flows, and, consequently, firm value. As pointed out by Peasnell et al. (2000: 420–421) the use of sub-optimal operating strategies is more costly than the reversals from accruals, and, consequently a more aggressive form of earnings management and thus, a last resource for management. However, even if real earnings manipulation is costly, bankruptcy is certainly even more so. In failed firms, the years prior to bankruptcy are probably characterised by more aggressive accounting practices, and in all likelihood, also by the undertaking of real actions to manage earnings and perceptions, particularly in those firms that have exhausted their possibilities for successful accrual manipulation. Therefore, we extend our analysis to study abnormal cash flow activity.

**H2:** Failed firms engage in real earnings manipulation in the years before failure. Real earnings management will be more pronounced if the firm's financial condition is weak and the possibilities for accrual manipulation have been exhausted.

*2.3.3. Effects of accounting and real earnings manipulation on the usefulness of earnings*

Prior research on earnings management in healthy firms shows that often discretionary accruals are value-relevant and are used by managers as a signalling device (Guay et al., 1996; Subramanyam, 1996). Healy and Palepu (1993) argue that income-increasing strategies have signalling value if they are used to communicate increases in future earnings to investors. In those cases, discretionary accruals can increase the usefulness of accounting numbers. However, as our sample comprises only ex-post bankrupt firms, we expect that income-increasing discretionary accruals in the years preceding failure will be due to opportunistic managerial behaviour.

Early work on failed firms only studies the magnitude of the proxies for income-increasing strategies to ascertain whether they respond to signalling vs. opportunistic reporting. We go one step beyond the descriptive evidence on the quantity of policy changes and discretionary accruals in Smith et al. (2001) and Rosner (2003) and explicitly analyse the consequences of managerial manipulation by studying if the reliability of accounting numbers in failed firms is hindered by their income increasing strategies. To do so, we study the conditional conservatism of earnings in failed vs. continuing firms.

We choose conditional conservatism as a benchmark to assess whether the usefulness of earnings declines in ex-post bankrupt firms as a result of both accounting and real activities manipulation for several reasons: (1) conditional conservatism is a summary measure of managerial reporting choices; (2) extant evidence demonstrates that better governed firms report more conditionally conservative earnings as a response to investors demands, who consider it as a desirable property of earnings (Beekes et al., 2004; Ahmed and Duellman, 2007; García Lara et al., 2007, 2009); and (3) recent research shows that conditionally conservative accounting leads to positive economic outcomes such as an ameliora-

[5] For example, if a covenant in a debt contract demands a minimum ROA of 2%, and true ROA is 1.9%, management may bridge the gap between both numbers by increasing net income artificially via accrual management that requires either accounting (e.g. reducing the provision for bad debts for the year) or real manipulation (e.g. relaxing credit terms to increase revenues), or both. This effect may also be achieved by selling assets, which would reduce the denominator (less assets) and increase the nominator (more cash or accruals), or by reducing discretionary expenses.

tion of the problems derived from information asymmetries (LaFond and Watts, 2008), or improvements in contracting efficiency (Ball et al., 2008).

Following Basu (1997), we define conditional conservatism as a timelier recognition of economic losses than economic gains in earnings.[6] For example, increases in the value of assets are not reflected in the profit and loss account, while decreases are reflected through impairments. At the root of this asymmetric treatment for the recognition of economic gains and losses in earnings is the principle of conservatism or prudence that is embedded into most regulatory accounting frameworks (FASB, 1980; IASC, 1989; ASB, 1999). This asymmetry in recognition requirements results in earnings that reflect bad news faster than good news (Basu, 1997). Timely loss recognition is expected to constrain management's opportunistic payments to themselves and to other parties, and is considered a desirable property of accounting earnings (Ball and Shivakumar, 2005) that mitigates moral hazard problems.

We hypothesise that in the years prior to failure, failed firms will show less conditionally conservative earnings than financially sound firms, because failed firms are likely to engage in less conservative accounting practices in those years to conceal their true economic performance (Ohlson, 1980), and in an attempt to postpone bankruptcy. As a consequence of this more aggressive (less conservative) accounting, managers relax the requirements for economic gains recognition, anticipating their recognition, while at the same time they attempt to delay the recognition of losses. Failing firms tend to have increasing inventories and increasing debtors (increase in accruals with reduced certainty of becoming cash flows). Although the ASB *Statement of Principles* (1999) does not prescribe conservative accounting, it states (ch. 3) that 'a degree of caution' has to be 'applied in exercising the necessary judgements'. Managers are thus expected to exercise caution and respond to this situation by increasing bad debt provisions and by impairing inventories. However, as pointed out by Ohlson (1980), when failure approaches, managers are expected to relax their conservative policies to try to delay it. Similar to Rosner (2003), we argue that if this is indeed the case, managers are likely to resort to aggressive accrual management in the years prior to failure. This aggressive accrual management could be achieved by, for example, not increasing bad debt provisions and not impairing inventories adequately.

Conditional conservatism is also expected to be affected by real earnings management. In an attempt to increase earnings, managers may aggressively relax credit terms, by selling on credit to customers with high ex-ante credit risk. This would lead to increases in the percentage of expected bad debts. Similar to the prior case, under the framework of the ASB, managers are expected to respond by increasing bad debt provisions. Failure to do so in an attempt to increase earnings would lead to aggressive instead of conservative accounting. The joint effect of both accruals and real activities manipulation in ex-post failed firms is expected to reduce the conservatism of accounting numbers in these firms. Consequently, our third and final hypothesis is as follows:

**H3:** Ex-post failed firms present less conditionally conservative earnings numbers than continuing firms in the years preceding bankruptcy.

## 3. Research method and sample selection procedure

We compare earnings quality between failed and continuing firms. Earnings quality is a broad concept with multiple dimensions (e.g. Al-Attar et al., 2008; Barker and Imam, 2008). We focus on earnings management and its influence on accounting reliability, as measured by conditional conservatism. Measures of earnings management and timely loss recognition are calculated for the full sample, to avoid biasing in favour of our hypothesis that earnings quality is lower in failed firms.[7] This section provides details of the calculation of these measures and describes the sample selection procedure.

### *3.1. Measurement of accounting accruals manipulation*

Ex-post failed firms may attempt to conceal deteriorating firm performance by using income-increasing accounting accruals. Extant research calculates abnormal accruals using the Jones (1991) model in cross-section, to improve the estimation of the parameters, as suggested by DeFond

[6] Beaver and Ryan (2005) identify two separate types of conservatism: conditional and unconditional. It is important to distinguish between them. They define unconditional conservatism as a persistent understatement of assets due to the use of accounting conventions at the inception of assets and liabilities, such as the non-recognition of internally generated intangible assets or the use of historical cost. Ball and Shivakumar (2005) argue that unconditional conservatism introduces a bias of unknown magnitude in the financial statements that does not contribute to contracting efficiency. Thus, in our study, we analyse only the conditional form.

[7] Prior research on failed firms like Rosner (2003) compares failed firms earnings quality estimates to those of matched non-bankrupt control samples, considering only firms with sound financial performance and avoiding cases that exhibit net losses. Excluding distressed firms biases the procedure in favour of finding the hypothesised result of differences between failed and continuing firms, especially if a number of continuing firms face financial problems, as happens in the UK (Neophytou and Mar Molinero, 2004).

and Jiambalvo (1994). The Jones model uses the unexplained part of a regression of total accruals on the change in revenue and gross property, plant and equipment as a proxy for abnormal accruals. We use the working capital version of the model because current research indicates that management has more discretion over current accruals, and that manipulation of long-term accruals such as depreciation is unlikely due to their high visibility and low flexibility (Becker et al., 1998; Young, 1999).

Aside from the working capital version of the Jones (1991) model, we also use the modified Jones model. The Jones (1991) model assumes that revenues are not discretionary. That is, the model disregards the possibility that managers also engage in real activities manipulation, by accruing revenues before the cash is received and when it is still questionable that the revenues have been earned. By assuming all revenues are non-discretionary, the Jones model removes part of the manipulation from the abnormal accrual proxy. Contrarily, the modified Jones model proposed by Dechow et al. (1995) assumes that all increases in credit sales are driven by managerial opportunistic decisions and classifies them as discretionary, modifying the Jones model by removing the change in receivables from the change in sales. Dechow et al. (1995) show that their modification ameliorates the tendency of the Jones model to measure abnormal accruals with error when discretion is exercised over revenues. Therefore, by using both the Jones and modified Jones models, we are able to analyse a proxy for manipulation capturing only 'pure' accounting manipulation (that is, the Jones model), and a proxy that pools together accounting and real activities manipulation (the modified Jones model). Finally, we employ the Kasznik (1999) model, which differs from the modified Jones model in that it incorporates the change in operating cash flow as an explanatory variable to account for the negative correlation between accruals and cash flows (Dechow, 1994).

To obtain a measure of abnormal working capital accruals for all firms in industry $j$ for year $t$, we estimate the Kasznik model cross-sectionally for all industry-year combinations with at least six observations of data, as follows:

$$\frac{WCA_t}{TA_{t-1}} = \alpha_0\left[\frac{1}{TA_{t-1}}\right] + \alpha_1\left[\frac{\Delta REV_t}{TA_{t-1}}\right] + \alpha_2\left[\frac{\Delta CFO_t}{TA_{t-1}}\right] + \varepsilon_t \qquad (1)$$

where, $WCA$ is working capital accruals, $\Delta REV$ is change in sales, $\Delta CFO$ is change in cash flow from operations and $TA$ are total assets, and $t$ is the time-period indicator. Next, for each firm, we calculate abnormal working capital accruals ($AWCA$) as follows:

$$AWCA_t = \frac{WCA_t}{TA_{t-1}} - \left(\hat{\alpha}_0\left[\frac{1}{TA_{t-1}}\right] + \hat{\alpha}_1\left[\frac{\Delta REV_t - \Delta REC_t}{TA_{t-1}}\right] + \hat{\alpha}_2\left[\frac{\Delta CFO_t}{TA_{t-1}}\right]\right) \qquad (2)$$

where, $\hat{\alpha}_0$, $\hat{\alpha}_1$ and $\hat{\alpha}_2$ are the fitted industry-coefficients from equation (1) and $\Delta REC$ is the change in accounts receivable. To run models (1) and (2), all available observations are used, including continuing and failed firms, to avoid introducing biases in the analysis. By removing the change in receivables ($\Delta REC$) from the change in revenues in equation (2) the model classifies as discretionary accruals all changes in receivables, including those driven by the manipulation of real activities.

To obtain abnormal accruals measures using the modified Jones model, we use the same procedure, but we do not include $\Delta CFO$ as an additional variable in models (1) and (2). Finally, to estimate abnormal accruals using the original Jones model, we follow this last procedure (i.e. exclude $\Delta CFO$ from the models), but without subtracting change in accounts receivable from change in sales in the second step. The Jones model measures discretionary accruals not including the effect of real activities manipulation, as it classifies all changes in accounts receivables as 'normal' accruals.

### *3.1.1. Classification of observations*

A key issue in our analysis of earnings manipulation by ex-post bankrupt firms is whether their attempts at masking poor performance in the years prior to failure were successful, and whether managers were able to mislead investors. To analyse if the accruals manipulation was successful at hiding poor performance we split sample observations in accordance to their ex-ante failure probability. Ex-post bankrupt firms that ex-ante do not show signs of failure are expected to have engaged in more aggressive accounting practices (Rosner, 2003), or at least, in practices that successfully masked poor performance.[8] To identify them, we calculate the ex-ante one-year-ahead probability of bankruptcy of all failed firms. We use the Charitou et al., (2004) failure prediction model, which is built using the logit methodology. The model is as follows:

$$P_{jt}(Y=1) = \frac{1}{(1+e^{-z})} \qquad (3a)$$

[8] An alternative explanation is that ex-post failed firms with large ex-ante failure probability do not have scope to manage accruals given their extremely deteriorated financial condition.

where,

$$-z = -7.1786 + 12.3826 * \frac{\text{Total Liabilities}_{jt}}{\text{Total Assets}_{jt}} - 20.9691 * \frac{\text{EBIT}_{jt}}{\text{Total Liabilities}_{jt}} - 3.0174 * \frac{\text{CFO}_{jt}}{\text{Total Liabilities}_{jt}} \quad (3b)$$

where, $P_{jt}$ $(Y = 1)$ is the probability of failure for entity $j$ at the end of year $t$; *EBIT* is earnings before interest and taxes; and *CFO* is cash flows from operations. Using this method, we obtain an ex-ante probability that the firm would end up being bankrupt that we use to identify firms that are suspect of more aggressive earnings practices.[9]

*3.2. Measurement of real earnings manipulation*

To analyse the existence of real activities manipulation we focus on one specific type of real earnings management: sales manipulation, which we measure following the method of Roychowdhury (2006). Similar to the calculation of abnormal accruals in equations (1) and (2), the first stage consists of deriving normal cash flow activity. We run the following cross-sectional regression for every industry-year combination with at least six observations of data:

$$\frac{CFO_t}{TA_{t-1}} = \beta_0\left[\frac{1}{TA_{t-1}}\right] + \beta_1\left[\frac{REV_t}{TA_{t-1}}\right] + \beta_2\left[\frac{\Delta REV_t}{TA_{t-1}}\right] + e_t \quad (4)$$

where all variables are defined as before. For every observation, abnormal cash flow (*ACFO*) is obtained by subtracting from actual firm *CFO* the normal *CFO* calculated using the estimated $\hat{\beta}$ coefficients from equation (4). The procedure is the same as the one previously explained for estimating a firm's abnormal accruals. Once *ACFO* has been estimated, we adapt the method by Roychowdhury (2006) to compare cash flow behaviour in suspect firms – in our case, the failed firms – with the sample of continuing firms as follows:

$$ACFO_t = \delta_0 + \delta_1 SIZE_{t-1} + \delta_2 NetIncome_t + \delta_3 FAILING_t + \tau_t \quad (5)$$

where *SIZE* is defined as the natural logarithm of total assets, *NetIncome* is net income scaled by beginning-of-period total assets, and *FAILING* is a dummy indicator taking the value of 1 if the firm goes bankrupt; 0 otherwise, and $t$ is the time-period indicator. If failed firms manipulate sales upwards by offering more lenient credit terms, then it is expected that the level of cash flow will be abnormally low, given the level of sales. Thus, if failed firms are manipulating their sales, $\delta_3$ will be negative and significant, consistent with an abnormally low level of *CFO*, given the reported sales.

*3.3. Measurement of conditional conservatism*

Both accounting accruals and real activities manipulation are likely to result in a reduction of timely loss recognition in failed firms. To measure the differential recognition speed of economic gains and losses in earnings we use the time series and accruals-based measures of conditional conservatism developed by Ball and Shivakumar (2005).

*3.3.1. Time series tests based on the persistence of income changes*

Basu (1997) and Ball and Shivakumar (2005) argue that the asymmetric recognition of economic gains and losses in earnings leads to differences in the persistence of gains and losses. Losses are less persistent as their timely recognition makes them appear as one-time shocks to earnings, while gains are reflected only when cash flows are realised, and, consequently, they are seen as permanent earnings components. This leads Basu (1997: 20) to argue that 'negative earnings changes have a greater tendency to reverse in the following period than positive earnings changes'. The tendency of negative earnings changes to reverse is documented by Brooks and Buckmaster (1976) and Elgers and Lo (1994).

Using the Basu (1997) method as a starting point,[10] Ball and Shivakumar (2005) develop a model to measure timely incorporation of gains and losses in accounting income 'as the tendency for increases and decreases in earnings to reverse' (p. 92). We follow Ball and Shivakumar (2005) and estimate the following model for our full sample:

$$\Delta NetIncome_t = \phi_0 + \phi_1 DNetIncome_{t-1} + \phi_2 \Delta NetIncome_{t-1} + \phi_3 DNetIncome_{t-1} * \Delta NetIncome_{t-1} + \sum \psi_t Year + \zeta_t \quad (6)$$

[9] We do not use Altman's (1968) Z-score model because it is built using LDA, a technique that makes strong demands on the structure of data. First, it requires that the financial ratios are normally distributed. This is known not to be the case (Ezzamel and Mar Molinero, 1987). Also, the ratios of failed companies should have the same variance-covariance structures as those of continuing firms. Richardson and Davidson (1983; 1984) show this assumption does not hold in the context of failure prediction. Furthermore, we do not use a model developed with US data for UK firms since there are significant differences between the reporting practices and insolvency codes of the two countries (Charitou et al., 2004). Finally, Mensah (1984) finds distress prediction models to be fundamentally unstable: coefficients varying according to the underlying health of the economy. Thus model derivation should be as close in time as possible to the period over which predictions are made.

[10] We do not use the Basu (1997) model as FAME does not provide share prices for most of our sample of bankrupt firms. In addition, several studies including Dietrich et al., (2007) recommend the use of non-market-based versions of Basu specifications, such as the ones in Ball and Shivakumar (2005).

where $\Delta NetIncome_t$ ($\Delta NetIncome_{t-1}$) is change in net income from fiscal year $t$-1 to $t$ ($t$-2 to $t$-1), scaled by beginning-of-period total assets. *DNetIncome* is a dummy variable that takes the value of 1 if the prior-year change in net income is negative; 0 otherwise. *Year* is a dummy variable for the fiscal year.

As described by Ball and Shivakumar (2005: 92), 'if gains are recognised in an untimely (smoothed) manner, they will be persistent components of net income that tend not to reverse and thus, $\phi_2$ will be equal to zero'. That is, under conservative accounting, they expect $\phi_2 = 0$. If managers choose aggressive accounting methods and foster good news recognition, positive income changes will behave as 'temporary earnings components that tend to reverse'. In our setting, where we compare ex-post failed with continuing firms, we expect that the $\phi_2$ coefficient will equal zero for continuing firms, as they will be conditionally conservative as required by accounting standards. However, for ex-post failed firms it will be significantly negative ($\phi_2 < 0$) if they engage, as expected, in income-increasing accounting practices to conceal poor performance and postpone bankruptcy. Regarding economic losses, their timely recognition implies that 'they are recognised as transitory income decreases, and hence reverse' (Ball and Shivakumar, 2005: 92). As they show, the implication is that $\phi_3$ will be negative. However, given that failed firms engage in aggressive accounting policies ($\phi_2 < 0$), in our setting we expect failed firms to show a significantly positive $\phi_3$ ($\phi_3 > 0$) if they delay the recognition of economic losses to the point that economic gains are reflected in earnings faster than economic losses. A positive $\phi_3$ implies economic losses are more persistent than economic gains. That is, firms deviate from conservative accounting and recognise losses in earnings at a slower pace than gains.

### *3.3.2. Accruals-based tests*

As developed in Ball and Shivakumar (2005), asymmetric gain and loss recognition results in an asymmetry in accruals, 'because economic losses are more likely to be recognised on a timely basis as unrealised (i.e. non-cash) accrued charges against income. Economic gains are more likely to be recognised when realised, and hence accounted for on a cash basis.' (p. 94). As a final test, following Ball and Shivakumar (2005) we estimate the model:

$$TACC_t = \gamma_0 + \gamma_1 DC_t + \gamma_2 CFO_t + \gamma_3 DC_t * CFO_t + \sum \psi_t Year + \upsilon_t \quad (7)$$

where *TACC* is total accruals divided by beginning of period total assets; *CFO* is cash flow from operations divided by beginning-of-period total assets; *DC* is a dummy variable that takes the value of 1 if *CFO* is negative, and 0 otherwise. *Year* is a dummy variable for the fiscal year. Given that Dechow (1994) and Dechow et al. (1998) document a negative relation between accruals and cash flows, and that this negative relation is embedded in the very same nature of the accounting process (accruals reverse), Ball and Shivakumar (2005) expect a negative $\gamma_2$, reflecting the negative association between accruals and cash flows. Ball and Shivakumar (2005: 94) predict a positive $\gamma_3$ under the hypothesis that accrued losses are more likely when the firm presents negative cash flows. That is, under conditional conservatism, in the case of economic losses, the decrease in earnings and the decrease in cash flows happen in the same period, and, consequently, the negative correlation between accruals and cash flows decrease in bad news periods. This is captured by a positive $\gamma_3$. Our prediction is that failed firms have lower quality earnings than continuing firms, and therefore, are less likely to recognise losses as transitory items. This implies lower asymmetry, and thus, $\gamma_3$ is expected to be lower for failed firms. In the case of aggressive accounting, $\gamma_3$ would be negative. If managers manipulate earnings upwards (by releasing negative accruals or decreasing cost of goods sold) this would increase the negative association between accruals and cash flows, even in bad news periods, contributing to a negative $\gamma_3$ (aggressive accounting). If managers engage in real earnings management, and increase sales (and therefore, debtors) but not cash flows, this would again translate into creating more positive accruals in periods of negative cash flows, again contributing to a negative association between accruals and cash flows even in bad news periods, that is, to a negative $\gamma_3$. Consequently, income-increasing strategies, both through accounting or real earnings manipulation, lead to a negative $\gamma_3$ coefficient.[11]

Accrual behaviour in years of positive cash flows is expected to be different in failed and continuing firms too. If managerial behaviour in failed firms results in accruals manipulation to increase earnings, this would translate into earnings capturing economic gains faster, and thus, the negative association between cash flows and accruals would be reduced. Consequently, we expect a neg-

[11] Exogenous economic events may contribute as well to this negative $\gamma_3$ coefficient. An example of these economic events would be if a major customer suddenly slows down payment without any sales management involved, leading to lower cash flows as compared to sales. The firm may decide not to make a provision against this slower payment. If the firm ran into a cash flow crisis, we would also observe a negative association between cash flows and accruals in the model, a symptom of aggressive accounting. Thus, the model captures cases of aggressive accounting, regardless of the origin and motive of the underlying economic events.

**Table 1**
**Sample selection procedure and distribution of observations**

**Panel A: Sample selection procedure**

| | | *Firm-years* | *Firms* |
|---|---|---|---|
| Firms with data available in FAME entering bankruptcy 1998–2004 | | 1,068 | 352 |
| Less: | Firms with less than three years of data previous to bankruptcy | (49) | (32) |
| | Finance, real estate and insurance firms | (63) | (21) |
| | Duplicated firms | (56) | (11) |
| | Firms with missing data to calculate abnormal accruals | (66) | (20) |
| Failed firms with data to run earnings management tests | | 834 | 268 |
| Continuing firms with full data available in FAME 1995–2004 | | 14,742 | 2,801 |

**Panel B: Distribution of failed firms by type of failure**

| | *Administration* | *Liquidation* | *Receivership* | *Total* |
|---|---|---|---|---|
| Number of firms | 77 | 110 | 81 | 268 |
| (proportion) | (28.73%) | (41.04%) | (30.23%) | |
| Number of firm-years | 234 | 331 | 269 | 834 |
| (proportion) | (28.06%) | (39.69%) | (32.25%) | |

*Notes*
Failed firms are obtained from the FAME 'inactive companies' file. *Abnormal Accruals* are abnormal working capital accruals as calculated using the Kasznik (1999) model.

ative coefficient on $\gamma_2$ for failed and continuing firms, but smaller in absolute value for failed firms.

*3.4. Sample selection procedure*

Financial statement data are collected from the FAME database. FAME contains detailed information on public (quoted at either the London Stock Exchange's Primary Market or at the Alternative Investment Market (AIM)) and private firms in the UK and Ireland. All British publicly quoted firms included in the database 'Active company file' at the time of data collection entered our continuing sample. Where available, we collected data for these firms as far back as 1990. The continuing sample with full data available consists of 2,801 firms.

UK publicly quoted firms that received an administration order, went into receivership, or were liquidated, as per the *Insolvency Act 1986*, were identified from the FAME database 'Inactive company file'.[12] Furthermore, for a failed firm to qualify for our sample, it had to declare insolvency between 1998 and 2004, and to have at least three years of full financial data prior to its collapse. We exclude finance, insurance and real estate firms (division H, 60–67, of US SIC code system), as firms operating in these industries are structurally different and their financial reporting practices generally preclude combining them with non-financial firms (Gilbert et al., 1990). We also exclude duplicates and observations with missing data to run our earnings management tests.

These criteria result in 268 failed firms with full data available. From those, 77 went into administration, 81 into receivership, and 110 were liquidations. We cannot distinguish between firms in the main market and AIM as once firms become inactive FAME classifies them as 'Quoted Inactive'. Table 1 summarises the sample selection procedure and provides a classification of observations across types of failure.[13]

## 4. Empirical results

### *4.1. Accruals manipulation in failed vs. continuing firms*

Using the method described in Section 3.1.1 above, we first classify failed firm-year observations that have an ex-ante probability of failure lower than 15% as having a low probability

[12] From the 'Inactive company' file we choose companies registered in the UK, and select 'public companies' ('legal form') including only receiverships, administrations and liquidations. To confirm the nature of the failure and its formal date we look at announcements from other sources (e.g. press announcements). The list of failed firms is available from the authors upon request.

[13] Auditors play a significant role in ensuring financial statement quality, reducing the incidence of earnings management (Becker et al., 1998) and opportunistic loss deferrals (Chung et al., 2003). We do not consider the monitoring role of auditors in our study due to insufficient data. Out of our sample of failed firms, 25 observations received a qualified opinion (19 of them in $t$-1), 277 an unqualified opinion and for all others data were not available. The year prior to bankruptcy, 10.1% of failed firms received a qualified opinion, while 39.7% received an unqualified opinion and no data were available for the remaining firms.

**Table 2**
**Descriptive statistics of sample firms**

**Panel A: Continuing firms (N=14,742)**

| | *Mean* | *Median* | *Std. Dev* | *Q1* | *Q3* | *Min.* | *Max.* |
|---|---|---|---|---|---|---|---|
| Size | 9.686 | 9.272 | 2.202 | 8.031 | 10.978 | 6.099 | 15.878 |
| Revenue | 2.098 | 1.650 | 3.005 | 0.958 | 2.685 | 0.054 | 8.676 |
| Net income | 0.026 | 0.043 | 0.779 | 0.001 | 0.097 | –0.715 | 0.487 |
| Cash flow | 0.071 | 0.092 | 0.654 | 0.002 | 0.180 | –0.932 | 0.675 |
| Total accruals | –0.045 | –0.055 | 0.606 | –0.119 | 0.013 | –0.551 | 0.694 |
| Working cap. accruals | 0.003 | –0.017 | 0.434 | –0.067 | 0.047 | –0.498 | 0.739 |
| Abnormal accruals | 0.003 | 0.001 | 0.167 | –0.037 | 0.041 | –0.334 | 0.333 |
| Failure probability | 0.361 | 0.127 | 0.403 | 0.004 | 0.821 | 0.000 | 1.000 |

**Panel B: HighP failed firms (N=612)**

| | *Mean* | *Median* | *Std. Dev* | *Q1* | *Q3* | *Min.* | *Max.* |
|---|---|---|---|---|---|---|---|
| Size | 8.777 | 8.516 | 1.536 | 7.665 | 9.553 | 6.441 | 14.314 |
| Revenue | 2.585 | 2.234 | 1.705 | 1.506 | 3.291 | 0.303 | 9.103 |
| Net income | –0.060 | 0.001 | 0.231 | –0.074 | 0.036 | –1.155 | 0.219 |
| Cash flow | 0.007 | 0.035 | 0.239 | –0.063 | 0.115 | –1.025 | 0.450 |
| Total accruals | –0.066 | –0.073 | 0.197 | –0.149 | 0.015 | –0.815 | 0.447 |
| Working cap. accruals | –0.014 | –0.023 | 0.192 | –0.093 | 0.063 | –0.539 | 0.529 |
| Abnormal accruals | –0.036 | –0.027 | 0.113 | –0.076 | 0.013 | –0.391 | 0.204 |
| Failure probability | 0.781 | 0.910 | 0.259 | 0.617 | 0.991 | 0.165 | 1.000 |

**Panel C: LowP failed firms (N=222)**

| | *Mean* | *Median* | *Std. Dev* | *Q1* | *Q3* | *Min.* | *Max.* |
|---|---|---|---|---|---|---|---|
| Size | 8.982 | 8.870 | 1.736 | 7.671 | 10.059 | 5.767 | 14.259 |
| Revenue | 2.619 | 1.946 | 2.111 | 1.285 | 3.531 | 0.073 | 11.756 |
| Net income | 0.071 | 0.063 | 0.204 | 0.014 | 0.119 | –0.526 | 0.707 |
| Cash flow | 0.123 | 0.117 | 0.283 | 0.012 | 0.211 | –0.803 | 1.094 |
| Total accruals | –0.051 | –0.068 | 0.211 | –0.140 | 0.013 | –0.475 | 0.657 |
| Working cap. accruals | –0.001 | –0.017 | 0.205 | –0.085 | 0.059 | –0.411 | 0.687 |
| Abnormal accruals | 0.017 | 0.001 | 0.118 | –0.052 | 0.066 | –0.309 | 0.394 |
| Failure probability | 0.030 | 0.009 | 0.040 | 0.001 | 0.051 | 0.000 | 0.136 |

*Notes*
*HighP (LowP)* are failed firms with a high (low) ex-ante probability of failure. A high (low) probability is set up as being higher or equal (lower) than 15%.
*Size* is the natural logarithm of total assets. All other variables are scaled by beginning-of-period total assets: Revenue is sales. *Working Cap Accruals* is working capital accruals. *Failure Probability* is the probability that a firm will go bankrupt calculated as in Charitou et al. (2004). *Abnormal Accruals* is a measure of firm abnormal working capital accruals calculated using the Kasznik (1999) model. Firms are classified as (1) *Continuing* if they do not go bankrupt during the considered period (1998–2004), (2) *Failed (LowP)* if they go bankrupt and their ex-ante bankruptcy probability of going bankrupt is below 15%; or (3) *Failed (HighP)* if they go bankrupt and their ex-ante bankruptcy probability of going bankrupt is equal or above 15%.

(*LowP*) of failure, and all others as having a high probability (*HighP*) of failure.[14]

Table 2 presents descriptive statistics of sample firms. Continuing firms have a median (mean) probability of bankruptcy of 12.7% (36.1%). Failed firms are split between *LowP* and *HighP* firm-year observations.[15] *HighP* firms have a median (mean) probability of bankruptcy of 91.0% (78.1%), whilst *LowP* firms have a probability of 1% (3%). A more detailed look at the descriptive statistics reveals that *LowP* firms have higher net income (*NI*) and cash from operations (*CFO*) than both continuing and *HighP* firms. Interestingly, *HighP* firms have, on average, higher sales relative

[14] Observations are classified almost identically if we increase the threshold of failure to 20% or even 25%.

[15] Throughout the text we make reference to *HighP* and *LowP* firms for simplicity. It would be more correct to use the terms *HighP* and *LowP* firm-year observations.

**Table 3**
**Descriptive statistics of failed firms by year of bankruptcy**

**Panel A: *t*-4**

| | *LowP failed firms* | | | *HighP failed firms* | | | *p-value difference* | |
|---|---|---|---|---|---|---|---|---|
| | Mean | Median | Std. Dev | Mean | Median | Std. Dev | Mean | Medians |
| Size | 8.654 | 8.137 | 1.438 | 9.471 | 9.416 | 1.034 | (0.18) | (0.13) |
| Revenue | 1.183 | 1.082 | 0.921 | 2.098 | 1.878 | 1.405 | (0.13) | (0.19) |
| Net income | 0.025 | 0.013 | 0.083 | –0.001 | 0.006 | 0.026 | (0.26) | (0.45) |
| Cash flow | 0.064 | 0.103 | 0.082 | 0.110 | 0.121 | 0.104 | (0.24) | (0.27) |
| Total accruals | –0.039 | –0.059 | 0.062 | –0.112 | –0.114 | 0.090 | (0.09) | (0.19) |
| Working cap. accruals | –0.007 | –0.037 | 0.064 | –0.040 | –0.038 | 0.039 | (0.20) | (0.27) |
| Failure probability | 0.014 | 0.003 | 0.023 | 0.612 | 0.606 | 0.259 | (0.00) | (0.01) |

**Panel B: *t*-3**

| | *LowP failed firms* | | | *HighP failed firms* | | | *p-value difference* | |
|---|---|---|---|---|---|---|---|---|
| | Mean | Median | Std. Dev | Mean | Median | Std. Dev | Mean | Medians |
| Size | 9.269 | 9.259 | 1.484 | 8.797 | 8.249 | 1.654 | (0.03) | (0.01) |
| Revenue | 2.675 | 1.958 | 2.184 | 2.823 | 2.228 | 1.953 | (0.33) | (0.10) |
| Net income | 0.098 | 0.073 | 0.234 | 0.001 | 0.011 | 0.084 | (0.00) | (0.01) |
| Cash flow | 0.174 | 0.134 | 0.304 | 0.030 | 0.024 | 0.134 | (0.00) | (0.01) |
| Total accruals | –0.076 | –0.079 | 0.164 | –0.029 | –0.032 | 0.140 | (0.03) | (0.02) |
| Working cap. accruals | –0.023 | –0.021 | 0.160 | 0.014 | 0.006 | 0.140 | (0.06) | (0.03) |
| Failure probability | 0.030 | 0.006 | 0.042 | 0.742 | 0.847 | 0.269 | (0.00) | (0.00) |

**Panel C: *t*-2**

| | *LowP failed firms* | | | *HighP failed firms* | | | *p-value difference* | |
|---|---|---|---|---|---|---|---|---|
| | Mean | Median | Std. Dev | Mean | Median | Std. Dev | Mean | Medians |
| Size | 9.070 | 8.964 | 1.804 | 8.923 | 8.673 | 1.558 | (0.28) | (0.26) |
| Revenue | 2.436 | 1.834 | 1.623 | 2.592 | 2.396 | 1.346 | (0.25) | (0.08) |
| Net income | 0.087 | 0.049 | 0.190 | –0.030 | 0.003 | 0.124 | (0.00) | (0.00) |
| Cash flow | 0.160 | 0.137 | 0.257 | 0.039 | 0.064 | 0.143 | (0.00) | (0.00) |
| Total accruals | –0.073 | –0.075 | 0.149 | –0.069 | –0.080 | 0.113 | (0.40) | (0.46) |
| Working cap. accruals | –0.024 | –0.021 | 0.152 | –0.020 | –0.030 | 0.115 | (0.41) | (0.47) |
| Failure probability | 0.033 | 0.010 | 0.044 | 0.742 | 0.907 | 0.288 | (0.00) | (0.00) |

**Panel D: *t*-1**

| | *LowP failed firms* | | | *HighP failed firms* | | | *p-value difference* | |
|---|---|---|---|---|---|---|---|---|
| | Mean | Median | Std. Dev | Mean | Median | Std. Dev | Mean | Medians |
| Size | 8.977 | 8.782 | 1.967 | 8.813 | 8.597 | 1.458 | (0.25) | (0.39) |
| Revenue | 2.272 | 1.636 | 1.781 | 2.376 | 1.999 | 1.659 | (0.35) | (0.14) |
| Net income | 0.015 | 0.021 | 0.209 | –0.106 | –0.040 | 0.192 | (0.00) | (0.00) |
| Cash flow | 0.074 | 0.093 | 0.258 | 0.011 | 0.036 | 0.196 | (0.03) | (0.00) |
| Total accruals | –0.059 | –0.101 | 0.292 | –0.117 | –0.100 | 0.169 | (0.04) | (0.19) |
| Working cap. accruals | –0.015 | –0.042 | 0.283 | –0.065 | –0.053 | 0.172 | (0.06) | (0.27) |
| Failure probability | 0.041 | 0.021 | 0.045 | 0.865 | 0.966 | 0.211 | (0.00) | (0.01) |

*Notes*
*HighP (LowP)* are failed firms with a high (low) ex-ante probability of failure. A high (low) probability is set up as being higher or equal (lower) than 15%. *Size* is the natural logarithm of total assets. *Revenue* is sales. *Working cap. accruals* is working capital accruals. *Failure probability* is the probability that a firm will go bankrupt calculated as in Charitou et al. (2004). *Abnormal accruals* is a measure of firm abnormal working capital accruals calculated using the Kasznik (1999) model. Firms are classified as (1) *Continuing* if they do not go bankrupt during the considered period (1998–2004), (2) *Failed (LowP)* if they go bankrupt and their ex-ante bankruptcy probability of going bankrupt is below 15%; or (3) *Failed (HighP)* if they go bankrupt and their ex-ante bankruptcy probability of going bankrupt is equal or above 15%.

**Table 4**
**Yearly differences in abnormal accruals**

**Panel A: *t*-4**

| | Expected sign of difference | *LowP vs. HighP* Mean (p-value) | Median (p-value) | *LowP vs. Cont* Mean (p-value) | Median (p-value) | *Failed vs. Cont* Mean (p-value) | Median (p-value) |
|---|---|---|---|---|---|---|---|
| Abnormal accruals | (+) | 0.07<br>(0.01) | 0.02<br>(0.00) | 0.02<br>(0.05) | 0.02<br>(0.07) | 0.01<br>(0.91) | 0.01<br>(0.92) |

**Panel B: *t*-3**

| | Expected sign of difference | *LowP vs. HighP* Mean (p-value) | Median (p-value) | *LowP vs. Cont* Mean (p-value) | Median (p-value) | *Failed vs. Cont* Mean (p-value) | Median (p-value) |
|---|---|---|---|---|---|---|---|
| Abnormal accruals | (+) | 0.05<br>(0.00) | 0.04<br>(0.01) | 0.02<br>(0.09) | 0.02<br>(0.05) | 0.00<br>(0.27) | –0.01<br>(0.10) |

**Panel C: *t*-2**

| | Expected sign of difference | *LowP vs. HighP* Mean (p-value) | Median (p-value) | *LowP vs. Cont* Mean (p-value) | Median (p-value) | *Failed vs. Cont* Mean (p-value) | Median (p-value) |
|---|---|---|---|---|---|---|---|
| Abnormal accruals | (+) | 0.04<br>(0.00) | 0.04<br>(0.01) | 0.03<br>(0.06) | 0.02<br>(0.41) | 0.00<br>(0.12) | –0.01<br>(0.19) |

**Panel D: *t*-1**

| | Expected sign of difference | *LowP vs. HighP* Mean (p-value) | Median (p-value) | *LowP vs. Cont* Mean (p-value) | Median (p-value) | *Failed vs. Cont* Mean (p-value) | Median (p-value) |
|---|---|---|---|---|---|---|---|
| Abnormal accruals | (–) | –0.05<br>(0.02) | –0.04<br>(0.00) | –0.05<br>(0.01) | –0.05<br>(0.02) | –0.04<br>(0.00) | –0.03<br>(0.00) |

*Notes*
Firms are classified as (1) Cont (*continuing*) if they do not go bankrupt during the period studied (1998–2004), (2) *LowP* if they go bankrupt and their ex-ante probability of failing is below 15%; or (3) *HighP* if they go bankrupt and their ex-ante probability of failing is equal to or above 15%. To compare failed and continuing firms, every failed firm-year is matched to the average abnormal accruals of continuing firms in the same industry and year.
Reported $p$-values are for a one-tail test of differences, calculated using a t-test for the means, and a Wilcoxon test for the medians.
*Abnormal accruals* are abnormal working capital accruals as measured by the Kasznik (1999) model.

to cash flow than both *LowP* and continuing firms. The difference in mean and median *CFO* between these groups is significant at the 1% level using a $t$- and a Wilcoxon-test, respectively, while the difference in median sales is significant at the 5% level ($p$-value=0.022). This initial evidence is consistent with the sales manipulation hypothesis. *LowP* firms have significantly higher AWCAs than *HighP* firms, suggesting that *LowP* firms are potentially more successful at hiding their distress from the market, and can therefore manage earnings more aggressively than the sub-set of *HighP* firms, which show more signs of being distressed. The difference in mean and median abnormal accruals between the *HighP* and *LowP* groups is significant at the 1% level, ($t$=4.28, and $Z$=4.61).

Tables 3 and 4 analyse the evolution of key financial indicators during the years prior to bankruptcy for *HighP* and *LowP* firms. Extant research on bankruptcy suggests that significant differences exist between the accounting numbers of failed and non-failed firms up to five years before bankruptcy (Beaver, 1966). However, we expect that some firms (those classified as *LowP*) retain sufficient scope to manipulate their numbers, and thus, may successfully appear as healthy in the years prior to bankruptcy. Such successful manipulation could affect the classification performance of bankruptcy models, especially for a great number of firms that lie in what Jones (1987) refers to as

the 'grey' area.

Table 3, Panels A to D show the evolution of some basic financial indicators of *LowP* and *HighP* firms from four years prior to bankruptcy (*t*-4) until the year prior to the event (*t*-1), where t is the year of bankruptcy. In year *t*-4 the only relevant difference between *LowP* and *HighP* firms is on their ex-ante probability of failure. However, as bankruptcy approaches (years *t*-3 to *t*-1), *LowP* firms manage to stay more profitable (higher net income, cash flows and less negative accruals) than *HighP* firms. This evidence could be consistent with *LowP* firms successfully manipulating their earnings upwards prior to bankruptcy.

Table 4 analyses the evolution of abnormal accruals (calculated with the Kasznik, 1999 model), comparing *LowP* with continuing and *HighP* firms, as well as continuing firms with all failed firms grouped together. To compare failed and continuing firms, every failed firm-year is matched to the average abnormal accruals of continuing firms in the same industry and year. *LowP* firms are expected to manage earnings more aggressively than *HighP* and continuing firms. If they manipulate their earnings successfully in the years preceding bankruptcy, it is expected that in the fourth, third and second years before the failure event (i.e. *t*-4 to *t*-2) the difference in mean (median) *AWCA* will be positive. However, because these firms end up being bankrupt, there should be a reversal of the previously manipulated accruals in the year just before bankruptcy (*t*-1), and thus, the difference in *AWCA* should become negative in this last year. In accordance with our expectations, Table 4, Panels A to C show that the differences between *LowP* and *HighP* are significantly positive from *t*-4 to *t*-2. This is consistent with *LowP* firms manipulating earnings upwards in the years prior to failure. Also in line with our predictions, we can observe that the difference becomes negative the year just before bankruptcy (*t*-1). The results of year *t*-1, with a larger reversal of accruals for *LowP* firms are consistent with larger income-increasing earnings management in prior years. These results also discard the alternative explanation for the existence of the relatively large number of *LowP* firms that these firms are not real failures but voluntary liquidations.

The differences in *AWCA* between *LowP* and continuing firms presented in the middle column of Table 4 follow a similar pattern, albeit the significance is slightly lower. Overall, the evidence supports our hypothesis that managers of distressed companies engage in earnings management to conceal their poor performance and, as bankruptcy looms closer and no improvement in the situation is observed, they eventually exhaust their instruments for successful accrual manipulation, and an accumulation of bad news will be observed.

The last column of Table 4 shows the difference in *AWCA* between all failed firms (*LowP* and *HighP* combined) and continuing firms. The differences are not significant except for *t*-1, where negative accruals accumulate both for *LowP* and *HighP* firms. Thus, whilst *LowP* firms show signs of having aggressive accounting policies in place, *HighP* firms are likely to have exhausted their resources for successful manipulation, having average negative abnormal accruals (as shown on Table 2), compensating those of the *LowP* firms when all failed firms are pooled together. These results might imply that *HighP* firms are being closely monitored by the market as obvious cases of distress companies, whilst the market might not monitor *LowP* firms as closely, permitting a higher degree of manipulation in these companies.

Our results are robust to the use of other models of discretionary accruals. Specifically, we use the modified Jones model (Dechow et al., 1995), and the original Jones (1991) model. The modified Jones model classifies all increases in debtors as discretionary accruals, consequently capturing both pure accounting manipulation and real activities manipulation. The Jones (1991) model captures only pure accounting manipulation, as it classifies all increases in debtors as normal accruals.

### *4.2. Real activities manipulation*

Table 5, Panel A reports parameter estimates of running equation (4) to identify abnormally low cash flow levels relative to sales. We obtain results consistent with previous research by Roychowdhury (2006).[16] Table 5, Panel B reports descriptive statistics of normal (*NCFO*) and abnormal cash flow from operations (*ACFO*). These descriptive statistics are also consistent with previous research. Table 5, Panel C reports results of running equation (5) for sample firms. In accordance with our predictions, $\delta_3$ is significantly negative, consistent with failed firms having, on average, negative abnormal cash flows, given their reported levels of sales. *HighP* firms have the highest median level of sales of all firms in the sample (see Table 2); however, these sales do not create the expected level of cash flows. In fact, cash flows from operations in failed firms are abnormally low.

These results, combined with those of the prior section, provide some evidence consistent with the

[16] Whilst the calculation of abnormal accruals is a standard procedure in the literature, the calculation of abnormal cash flows (*ACFO*) is not yet so. Thus, we provide the results of running equation (4) and descriptive statistics of *ACFO* and *NCFO*. *ACFO* are calculated for the full sample of data available (20,049 observations) to avoid biasing the estimates of normal and abnormal *CFO*. Equation (5) is run for the set of firms with full data to run our earnings management tests (15,654 observations, see Table 1 for details).

**Table 5**
**Abnormal cash flow (ACFO) activity in failed firms**

**Panel A: Estimation of ACFO. Model parameters**

| | Intercept | $\beta_1$ | $\beta_2$ | Adj. R-sq | N |
|---|---|---|---|---|---|
| Parameter | 0.0003 | 0.0711 | –0.0345 | 0.43 | 20,049 |
| *t*-stat | (0.12) | (3.09) | (–0.50) | | |

**Panel B: Abnormal and normal CFO**

| | Mean | Median | Std. Dev | Q1 | Q3 | N |
|---|---|---|---|---|---|---|
| Abnormal CFO | 0.006 | 0.022 | 0.198 | –0.069 | 0.103 | 20,049 |
| Normal CFO | 0.074 | 0.063 | 0.079 | 0.029 | 0.109 | 20,049 |

**Panel C: Comparison of failed (HighP) firms with the all other firms**

| | Intercept | $\delta_1$ | $\delta_2$ | $\delta_3$ | Adj. R-sq | N |
|---|---|---|---|---|---|---|
| Parameter | 0.072 | –0.001 | 0.012 | –0.083 | 0.23 | 15,654 |
| *t*-stat | (8.71) | (–1.36) | (2.43) | (–21.45) | | |

*Notes*
Panel A presents results of running equation (4) cross-sectionally:

$$\frac{CFO_t}{TA_{t-1}} = \beta_0\left[\frac{1}{TA_{t-1}}\right] + \beta_1\left[\frac{REV_t}{TA_{t-1}}\right] + \beta_2\left[\frac{\Delta REV_t}{TA_{t-1}}\right] + e_t$$

where, *CFO* is cash flow from operations, *TA* is total assets, and *REV* is total sales. There are 380 industry-year combinations from 1993 to 2003 with at least 6 observations per industry; the average industry-year group has 53 observations. This panel reports mean coefficients, t-stats and $R^2$ across all industry-years.
Panel B reports abnormal and normal CFO estimates. Abnormal CFO (ACFO) is obtained subtracting from actual CFO the expected or normal firm CFO, which is obtained using the industry specific coefficients derived from equation (4) above.
Panel C shows results of running equation (5) for sample firms

$$ACFO_t = \delta_0 + \delta_1 SIZE_{t-1} + \delta_2 NI_t + \delta_3 FAILING_t + \tau_t$$

where, *SIZE* is the natural logarithm of total assets, *NI* is net income scaled by beginning-of-period total assets, and *FAILING* is a dummy indicator that takes the value of 1 if the firm goes bankrupt, 0 otherwise. Reported t-statistics are White (1980) heteroskedasticity consistent.

existence of potential preferences and trade-offs in choosing earnings management instruments. We show that *HighP* firms, having probably exhausted their opportunities for successful accrual (purely accounting) manipulation, resort to real actions that lead to an abnormal relation between cash flows and sales. This abnormal relation could be partly attributable to increases in sales through extended credit terms, beyond what could be considered 'normal' in the industry.

### *4.3. Conditional conservatism in ex-post failed firms*

Results from tests of timely loss recognition using time-series tests (Table 6) and accruals-based tests (Table 7) are consistent with managers of failed firms engaging in pure accounting manipulation, and also in real activities manipulation to boost earnings. This translates into less conditionally conservative earnings.

Table 6 shows the results of running equation (6) for the sample of continuing firms, the whole sample of failed firms, and the three sub-samples of failed firms in years prior to failure (*t*-1, *t*-2 and *t*-3). The results show a distinct different behaviour between continuing and failed firms. As predicted by Ball and Shivakumar (2005), the $\phi_2$ coefficient of continuing firms, consistent with conditionally conservative accounting, is very close to zero and not significant. Timelier recognition of economic gains leads to a significantly negative $\phi_2$ coefficient. This is the case when we look at the failed firms. Also, if we look at the coefficient capturing the recognition of economic losses ($\phi_3$) we can see that, for continuing firms, the coefficient is, as expected, significantly negative showing the asymmetric nature of earnings and the more contemporaneous recognition of economic

**Table 6**
**Time series test of timeliness in loss recognition in continuing and failed firms**

| | Continuing total | Failed Total | Failed t-1 | Failed t-2 | Failed t-3 |
|---|---|---|---|---|---|
| Constant ($\phi_0$) | –0.03 | –0.05 | –0.05 | –0.05 | 0.04 |
| *(t-stat)* | *–8.37* | *–3.79* | *–1.62* | *–2.63* | *0.66* |
| DNetIncome$_{t-1}$ ($\phi_1$) | –0.04 | 0.02 | –0.02 | 0.03 | –0.11 |
| *(t-stat)* | *–1.80* | *1.22* | *–0.59* | *1.46* | *–1.67* |
| ΔNetIncome$_{t-1}$ ($\phi_2$) | 0.01 | –0.32 | –0.89 | –0.24 | –0.44 |
| *(t-stat)* | *0.44* | *–2.30* | *–2.64* | *–2.26* | *–2.27* |
| DN$_{t-1}$ ΔNI$_{t-1}$ ($\phi_3$) | –0.76 | 0.40 | 1.07 | 0.20 | –0.47 |
| *(t-stat)* | *–3.23* | *2.14* | *2.86* | *1.31* | *–0.76* |
| t-stat diff. $\phi_2$ cont vs failed | | 2.34 | 2.66 | 2.30 | 2.31 |
| t-stat diff. $\phi_3$ cont vs failed | | –3.86 | –4.14 | –3.42 | –0.28 |
| Chow test (F-stat) diff. $\phi_2$ cont vs failed (p-val) | | 0.01 | 0.00 | 0.21 | 0.21 |
| Chow test (F-stat) diff. $\phi_3$ cont vs failed (p-val) | | 0.00 | 0.00 | 0.01 | 0.94 |
| Adj. $R^2$ | 0.12 | 0.07 | 0.06 | 0.11 | 0.09 |
| N. Obs. | 14,227 | 340 | 151 | 133 | 56 |

*Notes*
We run equation (6):

$$\Delta NetIncome_t = \phi_0 + \phi_1 DNetIncome_{t-1} + \phi_2 \Delta NetIncome_{t-1} + \phi_3 DNetIncome_{t-1} * \Delta NetIncome_{t-1} + \sum \psi_t Year + \zeta_t$$

where $\Delta NI$ is change in net income from fiscal year t-1 to t, scaled by beginning-of-period total assets. *DN* is a dummy variable that takes the value of 1 if the prior-year change in net income is negative; 0 otherwise. Year is a dummy variable for the fiscal year.

t-statistics are White (1980) heteroskedasticity-consistent.
We estimate the statistical significance of the differences between coefficients from different regressions using two procedures:

a) As in Giner and Rees (2001), we use the following statistic (distributed as a student t): $\theta_1 - \theta_2$ divided by $\sqrt{\sigma_1^2 + \sigma_2^2}$
where $\theta_i$ is the estimated coefficient and $\sigma_i$ the standard error for variable i.
b) The more standard Chow (Wald) test. This specification is based on running a pooled regression with dummy variables to identify the different samples (in our case, failed vs continuing). As pointed out by Giner and Rees (2001 footnote 5) this approach assumes that the variance of the error across groups is the same.

losses than economic gains in earnings. However, there is again a shift in the sign of the coefficient for failed firms, consistent with more aggressive accounting policies when financial health is in trouble. The difference in the $\phi_2$ and $\phi_3$ coefficients between continuing firms and the different sets of failed firms is significant at conventional levels in the first and second year before failure when we calculate the significance of the difference between failed and continuing firms using the approach in Giner and Rees (2001). When we use the more standard (and more restrictive) Chow test, we obtain similar results.

When we analyse economic gains and losses recognition in earnings by looking at the relation between accruals and cash flows (Table 7), we can see that, as predicted by Dechow (1994), Dechow et al. (1998) and Ball and Shivakumar (2005), there is a negative association between accruals and cash flows for all firms (negative and significant $\gamma_2$). Consistent also with our hypothesis of failed firms using more aggressive accounting, the $\gamma_2$ coefficient is significantly larger (smaller in absolute value) in failed than in continuing firms. Finally, the $\gamma_3$ coefficient, capturing the asymmetry in the recognition of economic gains and losses in

**Table 7**
**Accruals-based test of loss recognition in continuing and failed firms**

| | Continuing total | Failed Total | Failed $t$-1 | Failed $t$-2 | Failed $t$-3 |
|---|---|---|---|---|---|
| Constant ($\gamma_0$) | 0.06 | –0.07 | –0.07 | –0.05 | –0.05 |
| *(t-stat)* | *5.16* | *–6.14* | *–3.70* | *–4.58* | *–2.68* |
| $DCFO_t$ ($\gamma_1$) | 0.15 | –0.03 | –0.08 | –0.01 | 0.02 |
| *(t-stat)* | *2.21* | *–0.74* | *–1.78* | *–0.20* | *0.52* |
| $CFO_t$ ($\gamma_2$) | –0.91 | –0.28 | –0.50 | –0.29 | –0.20 |
| *(t-stat)* | *–13.06* | *–4.21* | *–5.14* | *–4.70* | *–3.23* |
| $DCFO_t$ $CFO_t$ ($\gamma_3$) | 1.32 | –0.92 | –0.74 | –1.00 | –0.99 |
| *(t-stat)* | *4.01* | *–2.80* | *–2.25* | *–2.86* | *–1.63* |
| t-stat diff. $\gamma_2$ cont vs failed | | –6.68 | –3.47 | –6.60 | –8.09 |
| t-stat diff. $\gamma_3$ cont vs failed | | 4.65 | 4.19 | 3.96 | 3.65 |
| Chow test (F-stat) diff. $\gamma_2$ cont vs failed (p-val) | | 0.02 | 0.64 | 0.04 | 0.00 |
| Chow test (F-stat) diff. $\gamma_3$ cont vs failed (p-val) | | 0.00 | 0.00 | 0.00 | 0.00 |
| Adj. $R^2$ | 0.52 | 0.44 | 0.49 | 0.49 | 0.44 |
| N. Obs. | 14,227 | 340 | 151 | 133 | 56 |

*Notes*
We run equation (7):

$$TACC_t = \gamma_0 + \gamma_1 DC_t + \gamma_2 CFO_t + \gamma_3 DC_t * CFO_t + \sum \psi_t Year + \upsilon_t$$

where *TACC* is total accruals divided by beginning of period total assets; *CFO* is cash flow from operations divided by beginning-of-period total assets; *DC* is a dummy variable that takes the value of 1 if *CFO* is negative, and 0 otherwise. Year is a dummy variable for the fiscal year.
t-statistics are White (1980) heteroskedasticity-consistent.
We estimate the statistical significance of the differences between coefficients from different regressions using two procedures:

a) As in Giner and Rees (2001), we use the following statistic (distributed as a student t): $\theta_1 – \theta_2$ divided by $\sqrt{\sigma_1^2 + \sigma_2^2}$ where $\theta_i$ is the estimated coefficient and $\sigma_i$ the standard error for variable i.
b) The more standard Chow (Wald) test. This specification is based on running a pooled regression with dummy variables to identify the different samples (in our case, failed vs continuing). As pointed out by Giner and Rees (2001 footnote 5) this approach assumes that the variance of the error across groups is the same.

earnings is, consistent with conservative accounting, significantly positive for continuing firms. However, this is not the case for failed firms, where the coefficient is significantly negative. This negative $\gamma_3$ coefficient is consistent with a faster recognition of economic gains than economics losses, and with aggressive (instead of conservative) accounting.[17] These aggressive (instead of conservative) earnings may be the result of: (1) managers in failed firms using accruals to increase earnings (pure accounting manipulation, release of negative accruals, decrease of cost of goods sold), which decreases the predicted negative association between earnings and cash flows; and (2) managers in failed firms relaxing credit policies to increase sales without recognising the proper bad debt provisions (real activities manipulation together with improper accounting policies). The difference in the conservative measures between failed and continuing firms is always significant at

[17] Notice that the number of observations from Tables 6 and 7 differ slightly from the sample used for the earnings management tests due to data requirements to estimate models 6 (time series tests) and 7 (accruals-based tests). If we allow the sample to differ between Tables 6 and 7, we increase substantially the sample size to estimate model 7 (as we have more stringent data requirements for model 6, Table 6), obtaining up to 514 firm-year observations for failed firms. Results of tests using this larger sample do not differ significantly from those reported in Table 7.

least at a 1% level when using the Giner and Rees (2001) approach. The Chow test offers the same results, except for the good news coefficient in the year prior to failure, where the difference between failed and continuing firms is not significant.[18]

Overall, our results from Tables 6 and 7 show that managers of ex-post failed firms delay the recognition of economic losses and anticipate the recognition of economic gains in an attempt to delay bankruptcy.

## 5. Summary and conclusions

This paper analyses earnings quality in a large sample of UK bankrupt firms. We find that ex-post failed firms engage in earnings manipulation which may begin up to four years prior to failure. We argue that this manipulation responds to managers' attempts at concealing poor performance. We show that managers use two different tools to achieve this objective: purely accounting (accruals) manipulation and real activities manipulation.

Our study adds to the literature on the quality of earnings reported by ex-post failed firms in several ways: (1) we analyse two different aspects of earnings manipulation: accounting manipulation and real activity manipulation. We show managers resort to both, but use more aggressive real activities manipulation when their failure probability is high. This is consistent with arguments in prior literature that managers see real activities manipulation as more costly, and use it only when the available accounting discretion has been used up. (2) We also provide evidence on the timing of the manipulation, as we show it starts four years prior to failure, and unravels in the year just before failure; (3) we show that the manipulation is successful in hiding poor performance (firms with low ex-ante failure probability show aggressive accruals management), which highlights the need for better corporate governance provisions and enforcement of accounting standards, and opens the door for further research on failure prediction models that explicitly control for earnings management practices; (4) we demonstrate that ex-post failed firms report less conditionally conservative earnings, which implies increased agency problems; and finally, (5) we use a UK sample, where the insolvency code allows for a wider definition of bankruptcy, with different implications, than in the US (Franks et al., 1996; Bradbury, 2007).

Regarding our finding that failed firms with a low ex-ante failure probability manage earnings more aggressively, it is very likely that it is precisely the manipulation of earnings that determines the low probability of failure. This result has important regulatory implications, as it suggests that monitoring suspect firms (those with a larger failure probability) is not enough to prevent the manipulation in firms with financial difficulties. This result also suggests that bankruptcy prediction models may not properly account for managerial accounting choices. This is a potential implication of our work that generates a research question of interest: whether bankruptcy prediction models could improve by explicitly considering accounting manipulation. Finally, the evidence reported on the manipulation of real activities, although admittedly limited due to data constraints, is also of interest, as it raises the issue of sub-optimal decision-making to conceal poor performance. This may be of relevance to claimants and debtholders, as sub-optimal decision-making may reduce the liquidation value of some of the firm assets prior to liquidation.

With respect to ex-post failed firms reporting less conditionally conservative earnings in the years preceding bankruptcy, this is a consequence of both the manipulation of accounting accruals relative to the current level of sales, and the manipulation of real activities to increase current sales. We demonstrate that managers of ex-post bankrupt firms attempt to delay bankruptcy by recognising economic gains earlier and delaying the recognition of economic losses. Given that prior studies show that better governed firms present more conditionally conservative accounting numbers, our results may suggest that failed firms have weaker governance structures that open the door for the manipulation of real activities and to increases in operational risk. However, this is an issue that requires further research to obtain direct evidence. The absence of conditionally conservative accounting practices worsens contracting efficiency, increasing agency costs that affect both debt-holders (as managers avoid breaking debt covenants), and shareholders (expropriated by managers through compensation contracts that would not have rewarded additional bonuses to managers given the actual performance).

The evidence provided in the present study can be, consequently, useful for regulators, to develop and implement corporate governance provisions to prevent managers' opportunistic behaviour; for auditors, to understand better how managers exercise the discretion inherent in accounting standards to mask poor performance in financially troubled firms; and to other parties (such as analysts, creditors and researchers) who use accounting numbers to assess failure probability, default risk and the liquidation value of the firm.

[18] These results are robust to the use of working capital discretionary accruals as the dependent variable in the model. This is indicative of the use of aggressive accounting (pure accounting manipulation).

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# Selection bias and the Big Four premium: new evidence using Heckman and matching models


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**Abstract**—Many prior studies have found that large auditors charge significantly higher fees for statutory audit services, potentially resulting from higher audit quality and/or a lack of competition in the audit market. However, recent research using a Heckman two-step procedure attributes the large auditor premium to auditor selection bias. In this paper we examine the limitations of the Heckman model and estimate the large auditor (Big Four) premium using decomposition and matching methods on a large sample of UK private companies. Our analysis suggests that Heckman two-step estimates are highly sensitive to changes in sample and model specification, particularly the presence of a valid identifying variable. In contrast, the propensity score and portfolio matching methods we employ point to a persistent large auditor premium, consistent with the majority of previous studies. Conclusions of the premium vanishing when selection bias is controlled for therefore appear premature. Since the Heckman model is increasingly used in auditing and other areas of accounting research, our discussion and findings are likely to be of more general interest.



**Keywords:** audit fees; Big Four premium; matching estimators; propensity score matching; selection bias; two-step models


## 1. Introduction and background

In recent years, the competitiveness of the market for audit services has been the subject of considerable attention from the accounting profession, regulators and academic researchers. Among the main issues of concern is whether Big Four auditors command a premium when setting fees for statutory corporate audit services and, if so, whether the premium is symptomatic of a lack of competition in the audit market, or results from a higher quality product in competitive markets. In the UK, the (then) Department for Trade and Industry and the Financial Reporting Council (Oxera, 2006) estimated the Big Four premium at 18%; they concluded that it results from higher concentration and while auditor reputation is important to companies, some large UK firms have no effective choice of auditor due to significant barriers to entry. Furthermore, since the seminal contribution of Simunic (1980), a large number of studies, from a variety of markets and countries, find a premium using ordinary least squares (OLS) regression for large (Big Eight, Big Six and Big Four) auditors, for companies of various sizes (e.g. Pong and Whittington, 1994; Seetharaman et al., 2002; McMeeking et al., 2007; Clatworthy and Peel, 2007). A survey of the international empirical evidence (Moizer, 1997: 61) reports that 'the results point to a top tier fee premium of between 16 to 37%'; meanwhile, in a meta-analysis of 147 published audit fee studies, Hay et al. (2006: 176) find that 'the results on audit quality strongly support the observation that the Big 8/6/5/4 is associated with higher fees.'

Notwithstanding these findings, recent research has turned its attention to the important issue of the non-random selection of auditors and its impact on observed large auditor premiums. Some studies report that the premium paid to large auditors is larger than implied by OLS estimates when the Heckman model is employed to control for selection bias. In a study of UK listed companies, Ireland and Lennox (2002: 89) conclude that 'the large audit fee premium is more than twice as large when one controls for selection bias (53.4% compared to 19.2%)'. Moreover, based on a sample of UK companies over the period 1985 to 2002, McMeeking et al. (2006) find that the large auditor premium increases when selection bias is controlled for. Not all research using the correction for selection bias has produced the same results, however. In a recent examination of the pricing of audit services in UK local authorities, Giroux and


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Jones (2007) report that the selection bias correction makes little difference to inferences from OLS regressions. Furthermore, in a study of UK private companies, Chaney et al. (2004) fail to find a large auditor premium after they control for potential self-selection. Using OLS regression, they find a significant positive coefficient on a large auditor (Big Five) binary variable, but no premium when a two-step Heckman procedure to control for potential self-selection of auditors is used. Indeed, they conclude (2004: 67) that 'if big 5 auditees had chosen non-big 5 auditors, their audit fees would have been higher.' Similar findings were reported by the same authors for a sample of US listed firms (Chaney et al., 2005). These findings represent a very important development in the literature since they imply that many previous studies may have erroneously reported large auditor premiums – despite recent evidence (e.g. Blokdijk et al., 2006) that the Big Four provide higher quality audits.

The purpose of this paper is to present new evidence on the Big Four auditor premium and the effects of auditor selection for a large sample (36,674) of independent private UK firms by employing two-stage estimators and new decomposition and matching methods. Although the Heckman estimator represents an innovative correction for selection bias, it requires a valid additional identifying variable (instrument) for reliable implementation of the method; however, such variables are often extremely difficult to obtain in practice.[1] Since previous studies using the Heckman estimator do not focus on this important issue, this paper reports the sensitivity of estimates to the inclusion/exclusion of such an instrument. Furthermore, the Heckman estimator has been employed uncritically in the accounting literature to date, but analyses in other social science research suggest that it is highly sensitive to model specification, in contrast to OLS single-stage estimates (e.g. Hartman, 1991; Stolzenberg and Relles, 1997). This is particularly important because the Heckman model is increasingly used as a 'robustness test' for selection bias in accounting research in general and in the auditing literature in particular. If the model is not properly identified, this may lead to Heckman results lacking robustness due to severe collinearity problems (Little and Rubin, 1987; Puhani, 2000). Our analysis therefore focuses on these issues by illustrating the effects and sensitivity of the Heckman model to specification and sample variations and to the valid identification of the two-stage equations. An independent contemporaneous working paper by Francis and Lennox (2008) also examines the sensitivity of the Heckman model in estimating the large auditor premium for UK private firms. Although there is naturally some overlap between the two studies (and their results are of relevance to the current paper, so are discussed where appropriate), there are some important differences. In contrast to the present paper, the major aim of Francis and Lennox (2008) is to replicate the findings of Chaney et al. (2004) for a more recent time period, using the same models and sampling design as Chaney et al. (2004). The present paper does not aim to reproduce the results of Chaney et al. (2004), since it is based on a much larger cross sectional (rather than panel-based) sample of UK private firms and our models include additional important variables in order to obtain a more reliable and representative estimate of the Big Four premium. Importantly, our analysis also involves decomposition and matching methods not employed by Francis and Lennox (2008).

A further objective of the paper is to examine the significant methodological issue of how to estimate causal treatment effects in accounting and auditing research (the appointment of a Big Four auditor in our case) using OLS, selection and matching models. We conduct a comprehensive analysis of fee differences between Big Four and non-Big Four auditees using maximum likelihood (ML) and formal decomposition measures for selection effects not previously employed in the accounting and auditing literature. An increasingly popular approach adopted in the applied econometrics and finance literature involves matching procedures, particularly propensity score matching methods (e.g. Black and Smith, 2004; Simonsen and Skipper, 2006; Li and Prabhala, 2007), although these have yet to be employed in auditing research. Using these methods, we present new evidence on the large auditor premium using closely matched samples (in respect of size, risk and complexity) of companies audited by Big Four and non-Big Four auditors. In summary, our paper extends extant research by identifying potential weaknesses of the Heckman model and presenting alternative procedures to estimate the Big Four premium.

Our work contributes to the empirical literature on selection bias and the Big Four premium by using the largest sample of UK firms yet studied. The richness of our data set and large sample size allow us to subject this important issue to considerable scrutiny.[2] Our results suggest that two-step

[1] For instance, Neumayer (2003: 655) notes 'The problem is that such an exclusionary variable is frequently impossible to find'; moreover, Bryson et al. (2002: 9) state that 'the identification of a suitable instrument is often a significant practical obstacle to successful implementation.'

[2] At 36,674 observations, our sample is substantially larger than the largest (6,198 observations) in the meta-analysis of over 140 audit fee studies by Hay et al. (2006). Moreover, our dataset includes a more comprehensive set of variables than prior studies of selection bias and private UK company audit fees. In particular, the model reported by Chaney et al. (2004) excludes the number of subsidiaries and a second corporate size variable (sales), both of which have been found important in previous research.

corrections for selection bias in audit fee models are very sensitive to model specification (including the absence of an identifying variable), and to the sample used – findings consistent with empirical results in applications of two-step models in other fields (e.g. Winship and Mare, 1992; Stolzenberg and Relles, 1997; Leung and Yu, 2000). Using more consistent propensity score and pre-processed portfolio matching approaches, we conclude that the Big Four premium is still present after controlling for observable audit client characteristics and that models attributing the premium to unobservable characteristics should be treated with a high degree of caution. The Heckman procedure is becoming more widely used in accounting research such as studies of accruals quality (Doyle et al., 2007) and effects of disclosure on stock returns (Tucker, 2007). Accordingly, our results and methods are also likely to be of more general interest to accounting and business researchers. The remainder of the paper is organised as follows. In the next section, we outline general modelling issues and assumptions; Section 3 describes our variables and data, while our empirical results based on single stage, two-step and matching estimators follow in Section 4. The paper concludes in Section 5 with a summary, implications and suggestions for future research.

## 2. Modelling issues and the Big Four premium

*2.1. Evidence on the premium in prior literature*

To date, the auditing literature has advanced several non-independent reasons for large auditors charging higher fees, including the Big Four (formerly Big Eight, Big Six and Big 5) being associated with established reputations, higher quality audits, higher training costs, higher potential losses in the event of shareholder litigation ('deep pockets') and the occupation of a position of oligopoly in many audit markets (Moizer, 1997). Craswell et al. (1995) note that in competitive markets, the large auditor premium represents a return to Big Four investments in brand name reputation for higher quality audits. In the market for the largest multinational companies, however, smaller auditors, due to their lack of technical resources and geographical coverage, are unable to compete; hence such auditees are limited in choice to Big Four auditors only. For example, the Oxera report (2006: i) concludes there are significant barriers to entry in the sub-market for large UK quoted companies, 'including the high cost of entry, a long payback period for any potential investment, and significant business risks when competing against the incumbents (Big Four) in the market'.

Testing whether or not the auditee market is competitive (i.e. amongst the Big Four) for the largest companies, or subject to cartel pricing behaviour, is clearly difficult, since no realistic counterfactuals exist – with, for example, Big Four auditors accounting for 97.4% of the audits of the FTSE 350 in 2005 (Oxera, 2006). In the current paper we study UK private companies, where the market is *a priori* competitive in that Big Four concentration is relatively low (8.3% of audits in our sample) and where both Big Four and non-Big Four auditors are represented across a wide range of auditee size. In such a market any observed premium is more likely to be related to perceived or actual audit quality differentials than to a lack of competition.[3]

In our modelling of audit fees, we therefore assume a competitive market using the seminal audit fee framework of Simunic (1980) and developed by Pong and Whittington (1994). Simunic (1980) hypothesises that audit fees vary in association with audit production functions, loss exposure and audit quality (modelled with reference to auditee size, complexity, risk and auditor quality). Pong and Whittington (1994) posit that supply is related to auditors' cost functions and hence largely associated with the quantity of work/effort. Because of professional and statutory prescriptions for minimum audit standards, Pong and Whittington (1994) argue that the demand for audit is relatively inelastic. Furthermore, as noted by Simunic (1980: 170), in terms of product differentiation, the audit market is hedonic, i.e. differentiated audit products (quality) are not directly observed and 'the principal differentiation characteristic of the service is likely to be the identity of the supplier ... it is the Big Eight firms which enjoy visibility and brand name recognition among buyers.'

The UK private company audit market is an interesting context in which to test for the presence of a large auditor premium. In addition to the more competitive nature of the supply side of the audit market, there are economic arguments both for and against the prediction that a Big Four premium will be observed. As argued by Chaney et al. (2004), lower agency costs for private firms (which are more closely held), potentially less reliance on financial statements by outsiders and lower litigation risk for auditors (compared to listed firms) would point to lower demand for high quality audit services, and hence to no expectation of a premium. By contrast, owners of private firms may seek to signal credibility of their financial

[3] Several studies indicate that a premium may be warranted as a result of differential audit quality. For example, Blokdijk et al. (2006) find that the quality of audits by the (then) Big Five is higher, even though the total effort exerted is similar to smaller auditors. Francis et al. (1999) report that Big Six auditors constrain income-increasing discretionary accruals more than smaller auditors, while Lennox (1999) finds that large auditors' reports are more accurate than those of smaller auditors.

**Figure 1**
**Linear regression for audit fees**

$$\ln F_{BIG4} = \alpha + \beta + \sum_{k=1}^{K} \beta_k X_{k\,BIG4} + \varepsilon_{BIG4} \qquad \text{For Big Four clients } (D = 1) \qquad (1)$$

$$\ln F_{NON} = \alpha + \beta + \sum_{k=1}^{K} \beta_k X_{k\,NON} + \varepsilon_{NON} \qquad \text{For non-Big Four clients } (D = 0) \qquad (2)$$

statements should they plan to sell their stake and the absence of market values may make information provided by the financial reporting process more important (e.g. for managerial performance measures). Collis et al. (2004) find increased demand for audit services in general (rather than between types of auditor) from firms wishing to maintain good relationships with lenders; it is possible that such findings may extend to firms wishing to appoint higher quality auditors to testify to the truth and fairness of their financial statements. In addition, there is evidence that newly listed firms attract cheaper debt capital if they appoint a large auditor (Pittman and Fortin, 2004), suggesting that higher audit fees may eventually be recovered through the payment of lower rates of interest.

*2.2. Statistical specifications and assumptions*

In this section, we describe the models and assumptions that provide the basis for our empirical analysis. Although our focus is on the Big Four premium, the discussion is applicable to other areas of accounting and business research where selection bias is a potential problem. At various points, we refer readers to the Appendix for a more formal exposition of the issues in this section.

*OLS and Oaxaca-Blinder decomposition*

We start by dividing companies into those with a Big Four auditor and those without. This division is indexed below by *BIG4* and *NON* and represented by a dummy variable ($D$) taking the value of one if the auditee appoints a Big Four auditor and zero otherwise. The literature typically assumes audit fees ($F$), expressed in natural log form ($\ln F$), depend on $K$ variables ($X_k$ $k = 1,..,K$) – principally auditee size, complexity and risk measures – employing a linear regression of the form shown in Figure 1, where the error term ($\varepsilon$) reflects unobservable random determinants of the fees paid to auditors.

Audit fees may vary between these groups because observable characteristics ($X$) are different and/or because the impact of these characteristics on audit fees ($\beta \neq 0$, $\alpha_k \neq \beta_k$) is different. For instance, as Pong and Whittington (1994) and Chaney et al. (2004) note, it is likely that Big Four auditors are better equipped to audit larger, more complex clients, although such comparative advantages may be offset in part by higher fixed costs associated with the training of audit staff.

Initially we assume for our single stage conventional estimates that any unobservable auditee characteristics are the same for $D = 1$ and $D = 0$, so the errors have the same distribution for each type of auditor. A problem arises since we cannot directly compare the fees paid under each regime because we only observe a company as a client of either a Big Four or a non-Big Four auditor, but not both, i.e. we do not observe the counterfactual outcome.[4] This problem can be overcome by assuming that the values of the regressors are unimportant in respect of computing the counterfactuals; however, if there are large and significant differences in the values of the regressors for Big Four and non-Big Four auditees, then it may be unreasonable to extrapolate between them.[5]

If the OLS estimates of the parameters in (1) and (2) are ($a$, $a_k$) for the non-Big Four auditees and ($a$, $b$, $b_k$) for the Big Four auditees, then the predicted log of audit fees for a Big Four auditee, firm $i$, in each audit regime are:

$$\widehat{\ln F_{NONi}} = a + \sum_{k=1}^{K} a_k X_{ki}$$

(the counterfactual value) and

$$\widehat{\ln F_{BIG4i}} = a + b + \sum_{k=1}^{K} b_k X_{ki}$$

(the actual predicted value). The Big Four premium

[4] Another potential concern is the use of linear functions. It may be possible for the same non-linear audit fee equation to apply to both types of auditee so that any observed Big Four premium might be entirely 'explained' by auditees' different characteristics. A premium can still be predicted if linear approximations are estimated at markedly different points on the curve.

[5] For example, at the limit, it would appear inappropriate to compare the audit fees paid by large and small auditees if *all* large auditees employed Big Four auditors while *all* small auditees employed non-Big Four auditors.

**Figure 2**
**Estimates of the Big Four premium**

$$P_{BIG4} = b + \sum_{k=1}^{K}(b_k - a_k)\overline{X_{kBIG4}} \quad (3)$$

$$P_{NON} = b + \sum_{k=1}^{K}(b_k - a_k)\overline{X_{kNON}} \quad (4)$$

is then the difference:

$$\widehat{\ln F_{BIG4i}} - \widehat{\ln F_{NONi}} = b + \sum_{k=1}^{K}(b_k - a_k)X_{ki}$$

(most previous studies test for a premium using a binary variable in a single regression, so the premium is constant at $b$). In practice we compute these statistics for two 'typical' (average) auditees: the first has the values for the regressors equal to the mean values for the Big Four auditees ($\overline{X_{kBIG4}}$) and the other the mean values for the non-Big Four auditees ($\overline{X_{kNON}}$).[6] This gives two estimates ($P$) of the Big Four premium, shown in Figure 2.

$P_{BIG4}$ represents the predicted fees paid by a typical Big Four auditee to a Big Four auditor minus the predicted log of fees paid by the same auditee to a non-Big Four auditor; whereas $P_{NON}$ represents the difference between predicted fees for a typical non-Big Four auditee paid to a Big Four auditor and the predicted fees paid to a non-Big Four auditor. Although not typically used in auditing research, these statistics are widely used elsewhere as part of an Oaxaca-Blinder (OB) decomposition analysis (see the Appendix, Oaxaca (1973), Blinder (1973) and Greene (2003: 53) for further details).

Recent developments in the auditing literature, however, point out that conventional OLS estimates of the Big Four premium are potentially biased since auditors are not appointed randomly by their clients and because auditor choice may be systematically related to auditees' unobservable characteristics (e.g. insiders' knowledge of the riskiness of future cash flows). Ireland and Lennox (2002: 75) note 'although the standard OLS audit fee models control for observable differences, characteristics that are not observable to the academic researcher may affect both fees and auditor choice and thereby cause bias.' However, it is not entirely clear from previous research what such unobservable characteristics represent or how important they are in systematically influencing auditor selection and audit fees. Titman and Trueman (1986) and Datar et al. (1991) each develop models predicting that auditor quality is a function of firm-specific risk, of which firm insiders are better informed than outsiders. But both models make competing predictions about the nature of the relationship between firm-specific risk and auditor quality: Datar et al. (1991) predict that entrepreneurs of risky firms choose higher quality auditors, whereas Titman and Trueman (1986) predict the opposite.

*Selection models*

Selection bias arises if the unobservable characteristics of Big Four and non-Big Four auditees are systematically different from each other. Suppose that $\varepsilon_{NON}$ and $\varepsilon_{BIG4}$ in equations (1) and (2) above are drawn from the same distribution but that Big Four auditees and non-Big Four auditees only have positive and negative errors respectively.[7] Estimating fee equations with standard single-stage OLS omits the conditional means (by assuming $E(\varepsilon_{BIG4}) = E(\varepsilon_{NON}) = 0$) and leads to inconsistent estimates if these terms are correlated with the regressors. The Heckman two-step procedure provides an estimate of the mean of the conditional error known as the Inverse Mills Ratio (IMR) or the selection term $\lambda$, which augments the regressors in equations (1) and (2) above. The procedure involves estimating a probit model of auditor choice as the first stage; this model yields estimates of selection terms $\lambda_{BIG4}$ and $\lambda_{NON}$ which are then included in the audit fee equations in the second stage. OLS applied to the augmented equations yields consistent coefficient estimates and standard hypothesis tests can be applied with modified formulae for the standard errors. The Heckman procedure thus estimates the equations shown in Figure 3.

The Appendix provides a more formal description of the model and the derivation of the selection terms. The probit and audit fee equations in

[6] This choice ensures that the errors play no role as the means of the predicted errors are zero and would seem reasonable on the basis that the mean represents the expected value of the characteristics of a Big Four auditee.

[7] For example, if the positive error measures the unobserved value to the auditee of appointing a Big Four auditor, then Big Four auditees will value Big Four auditors more than non-Big Four auditors and therefore pay higher fees.

**Figure 3**
**Equations for Heckman procedure**

$$\ln F_{BIG4} = \alpha + \beta + \sum_{k=1}^{K} \beta_k X_{k\,BIG4} + \sigma_{SBIG4}\lambda_{BIG4} + \upsilon \qquad \text{For Big Four auditees} \qquad (5)$$

$$\ln F_{NON} = \alpha + \sum_{k=1}^{K} \alpha_k X_{k\,NON} - \sigma_{SNON}\lambda_{NON} + \upsilon \qquad \text{For non-Big Four auditees} \qquad (6)$$

the Heckman model can also be estimated simultaneously by maximum likelihood (ML), which leads to more efficient estimates if the model is correctly specified. Accordingly, we report both conventional Heckman two-step and ML estimates in our empirical analysis.

Although the Heckman procedure has become increasingly popular in accounting research and in the finance literature (e.g. Li and Prabhala, 2007), its robustness has been questioned under certain conditions. For example, Giles (2003: 1299) notes 'Heckman's sample selectivity correction methodology offers a way of improving on the estimates obtained with non-random samples. While there is improvement in general in this regard, there are situations in which the correction for sample selectivity actually aggravates the problem.'

It is commonplace to assume joint normality of the distribution of the errors in the selection and outcome equations and that any systematic unobservable variables are normally distributed (an untestable assumption). Joint normality has the surprising and unfortunate implication that collinearity between the selection term and the other regressors in the second stage equation is often severe, leading to serious model instability (e.g. Leung and Yu, 2000). In addition, it is common for researchers to identify the second stage equation via the non-linearity of the selection term only. However, recent econometric analyses of this issue suggest that to adequately identify the model it should contain a valid instrument, i.e. a regressor which determines the choice of auditor but has no significant effect on determining audit fees (Little, 1985; and Puhani, 2000). But collinearity may cause problems even when an instrument (also known as an exclusion or identifying variable) is employed, leading to unstable estimates of treatment effects (Stolzenberg and Relles, 1997; Leung and Yu, 2000; Li and Prabhala, 2007). Against this background, it is perhaps unsurprising that, as noted above, empirical results in auditing research using the Heckman model have, to date, been mixed.[8]

In the absence of satisfactory instruments, the selection effect is best identified by extreme observations of the selection term $\lambda$, i.e. those companies with an estimated probability of choosing a Big Four auditor close to 1. These Big Four auditees (usually because of their large size and complexity) effectively have no surrogate non-Big Four counterfactuals – that is, there is no 'common support' – the common support region being where Big Four clients have non-Big Four counterparts with similar characteristics. Following Black and Smith (2004), we therefore assess the robustness of the Heckman results by estimating models using samples with different values of the selection term.

*Matching estimators*

The problems of model sensitivity, lack of robustness, linear functional form assumptions and the need for adequate counterfactuals motivate the use of matching methods. These methods are gaining in popularity in the applied econometrics literature (e.g. Bryson et al., 2002; Black and Smith, 2004; Simonsen and Skipper, 2006) and are based on matching the observable characteristics of members in the treatment group (i.e. Big Four auditees in our case) to members (counterfactuals) in the untreated group (non-Big Four auditees). Matching analyses are based on two important assumptions: (1) the conditional independence assumption (CIA) and (2) common support. The former requires the value of audit fees to be independent of auditor type given the values of some observable variables, whereas the latter involves comparable observations existing in both groups. Both assumptions are discussed at greater length in the Appendix.

A limitation of matching methods is that they cannot accommodate any systematic effects of unobservable auditee characteristics that have a joint impact upon auditor selection and audit fees (although as discussed above, it is not obvious what these systematic effects might be or what their directional influence is). Whereas the Heckman

[8] Though it is, of course, plausible that the variation in these findings is attributable to differences in the underlying economic situations (e.g. private limited versus public quoted company markets).

approach allows such unobservable factors to influence auditor choice systematically, it is often sensitive to specification and collinearity, as discussed above. Although they cannot deal directly with unobservables, matching estimators do not rely on linear extrapolation (outside the common support region) or functional form assumptions; nor do they require an exclusion variable or impose joint normality assumptions. As noted by Simonsen and Skipper (2006), matching methods are based on matching on observable characteristics of members in the treatment group (Big Four auditees in our research) to members (counterfactuals) in the untreated group (non-Big Four auditees) and hence 'balancing the bias arising from self-selection ... Matching allows for heterogeneous treatment effects, is not subject to parametric assumptions and does not *per se* assume separability of observables and unobservables' (*ibid.*: 920).

When applying matching methods, there are various estimators to choose from, reflecting a trade-off in respect of the number of variables used to match on, the closeness with which the variables are matched (particularly continuous variables), and the sample size. This is referred to as the 'curse of dimensionality' (Ho et al., 2007) since close matching on more than a few variables (dimensions) might result in matched samples that are too small for any meaningful analysis.

The first matching method we use – propensity score matching – has recently been employed in applied econometrics research (e.g. Black and Smith, 2004; Simonsen and Skipper, 2006), but not in the auditing literature to date. The method (see Appendix for more details) is implemented as follows. First the selection equation is estimated using a parametric estimator of the auditor selection equation (in our case a probit model) and the probabilities (propensity scores) of choosing a Big Four auditor are obtained for all sample firms. Each Big Four auditee is then matched to a non-Big Four auditee with a similar propensity score and differences in audit fees compared across the two matched samples. As noted by Black and Smith (2004: 110), the logic underpinning this method is that 'subgroups with values of X [explanatory variables] that imply the same probability of treatment can be combined because they will always appear in the treatment and (matched) comparison groups in the same proportion. As a result, any differences between subgroups with different X but the same propensity score balance out when constructing the estimates.' Hence, an important practical advantage of propensity score matching is that subgroups are matched on one variable, obviating the need for very large samples when subgroups are to be matched according to several characteristics. Moreover, propensity score matching does not make the same functional form assumptions as linear regression and non-linear relationships (of which there is some evidence in the literature – e.g. Peel and Roberts, 2003; Chaney et al., 2005) can be allowed for.

The second matching approach we take is an intermediate (semi-parametric) one which combines matching with the standard OLS regression used in the majority of prior studies. This approach involves matching observations according to important *actual* client characteristics prior to a standard parametric analysis, in line with the methods advocated by Ho et al. (2007), who highlight the potentially serious pitfalls of drawing inferences from sensitive statistical models. Initially we preprocess our data, then estimate the standard audit-fee model with a binary Big Four indicator variable. Preprocessing involves matching Big Four and non-Big Four auditees only on key attributes (well-tested measures of auditee size, complexity and risk) thereby ensuring a sufficient number of matched observations to conduct standard OLS regression techniques to control for any remaining confounding factors. We initially sort our full sample into quantiles based on the key attributes of sales (40 quantiles), the ratio of exports to sales (11 quantiles), return on total assets (40 quantiles) and the number of subsidiaries (10 quantiles). We then match each Big Four client to a non-Big Four client jointly sharing membership of each respective quantile for size, risk and complexity. It is important to note that this is a simultaneous requirement, i.e. matched auditees have similar size *and* risk *and* complexity characteristics and is therefore a stricter set of criteria than propensity score matching, since the latter is a composite (albeit conditional) score.

Since this process results in a large number of potential matched combinations, we perform this procedure 2,000 times. Each time, we estimate an OLS regression and for each iteration, we capture the coefficient for the binary Big Four indicator variable (representing the premium charged to similar Big Four and non-Big Four auditees) and report the results for the distribution of this coefficient. As stated by Ho et al. (2007: 3) this pre-processing approach combines the merits of both non-parametric matching with conventional parametric estimators: 'In a sense our recommendations already constitute best practice since matching alone is not a method of estimation and always requires some technique to compute estimates ... we simply point out that, except in the extraordinary case where matching is exact, parametric procedures have the potential to greatly improve causal inferences even after matching.' In summary, while the Heckman model presents a potential solution to the important problem of selection bias, it may produce imprecise results.

**Table 1**
**Variable definitions**

| *Label* | *Definition* |
|---|---|
| ln*AFEE* | Natural log of audit fee (in £) |
| ln*SAL* | Natural log of turnover (in £) |
| ln*TA* | Natural log of total assets (in £) |
| *SQSUBS* | Square root of the number of subsidiaries |
| *EXPSAL* | Ratio of non-UK turnover to total turnover |
| *QUALIF* | Binary variable taking the value of 1 if company had qualified audit report, 0 otherwise |
| *PBAL* | Binary variable taking the value of 1 if company disclosed a post-balance sheet event in accounts, 0 otherwise |
| *CONLIAB* | Binary variable taking the value of 1 if company disclosed contingent liabilities in accounts, 0 otherwise |
| *EXITEM* | Binary variable taking the value of 1 if company disclosed exceptional and/or extraordinary items in accounts, 0 otherwise |
| *RTA* | Ratio of profit before tax to total assets |
| *TLTA* | Ratio of total liabilities to total assets |
| *LOND* | Binary variable taking the value of 1 if company is located in London, 0 otherwise |
| *BUSY* | Binary variable taking the value of 1 if firm's year-end is in December or March, 0 otherwise |
| *BIG4* | Binary variable taking the value of 1 if company is audited by a Big Four auditor, 0 otherwise |
| *CHTA* | Absolute value of change in total assets from year $t$-1 to year $t$ |

Matching techniques are becoming increasingly popular, since as noted by Li and Prabhala, (2007: 51), 'They represent an attractive means of inference because they are simple to implement and yield readily interpretable estimates of "treatment effects"'. However, they are based on different assumptions to the Heckman model, principally because they assume that any unobservables are unimportant (Li and Prabhala, 2007). If unobservable client characteristics determine both auditor choice *and* audit fees, matching estimators produce potentially biased results. Since the theoretical research into the determinants of auditor choice is inconclusive and the nature and role of unobservables are unclear, however, matching methods seem an appropriate means to assess the robustness of the Big Four premium in recent auditing research.[9] The next section outlines the variables and data used in our empirical analysis.

## 3. Variables and data

### *3.1. Variables*

Our main empirical model of audit fees takes the standard linear form as in Figure 4.

The variables used in the model are described in Table 1 and have been widely employed in prior research (e.g. Simunic, 1980; Pong and Whittington, 1994; Chan et al., 1993; Ezzamel et al. 1996; Chaney et al., 2004; McMeeking et al., 2006; Clatworthy and Peel, 2007).[10]

Since corporate size (serving as a proxy for audit effort) has been found to be the key driver of external audit fees in previous research, we employ both total assets (£) and turnover (£) as auditee size measures in our research. Pong and Whittington (1994: 1075) note that audits have two broad dimensions: 'an audit of transactions and verification of assets. The former will be related to turnover and the latter to total assets.'

[9] Since prior research on this specific issue is relatively rare and has produced inconsistent results, further research into the identification and examination of such unobservable characteristics seems warranted.

[10] Naturally, the binary variable *BIG4* is omitted where the equations are estimated separately. As noted by an anonymous referee, a potentially important variable not included in our analysis is non-audit (consultancy) fees. This variable has been found to be significantly related to audit fees in a number of studies of listed companies; however, since private companies are not required to disclose their non-audit fees, the data are not available for most private companies (i.e. other than for those that voluntarily disclose them). This is a potential limitation of our research, since the distinction between the two types of fees is not always clear – though previous Heckman (two-step) research into UK private firms also omits non-audit fees, so our results are comparable in this respect.

**Figure 4**
**Empirical model of audit fees**

$$\ln AFEE = \alpha_0 + \beta_1 \ln SAL + \beta_2 \ln TA + \beta_3 SQSUBS + \beta_4 EXPSAL + \beta_5 QUALIF + \beta_6 PBAL + \beta_7 CONLIAB \quad (7)$$

$$+ \beta_8 EXITEM + \beta_9 RTA + \beta_{10} TLTA + \beta_{11} LOND + \beta_{12} BUSY + \beta_{13} BIG4 + \varepsilon$$

Following previous studies, we specify the relationship between audit fees (ln*AFEE*) and the size measures for turnover (ln*SAL*) and total assets (ln*TA*) in natural logarithmic form to capture potential economies of scale in the audit. In order to control for audit complexity, we include a variable labelled *SQSUBS*, defined as the square root of the number of subsidiaries (e.g. Francis and Simon, 1987), and *EXPSAL* – the ratio of non-UK turnover to total turnover (e.g. Beatty, 1993: Chaney et al., 2004), both of which we expect to be positively related to audit fees.

To capture auditee risk characteristics, we employ the gearing ratio of total liabilities to total assets (*TLTA*) and the ratio of net profit before tax to total assets (*RTA*), which we expect to be positively and negatively related to audit fees, respectively (e.g. Chan et al., 1993; Firth, 1997). Following previous research (e.g. Chaney et al., 2004; Clatworthy and Peel, 2007) we employ three additional binary variables to capture incremental risk/complexity in the audit. These are whether (coded 1) or not (coded 0) the audit client received a qualified audit report (*QUALIF*), reported exceptional and/or extraordinary items (*EXITEM*), disclosed a post-balance sheet event (*PBAL*) or a contingent liability (*CONLIAB*). All these variables are expected to be positively related to audit fees (*ibid.*).[11] Finally, we include binary variables for whether (coded 1) or not (coded 0) companies are audited by a Big Four auditor (*BIG4*), whether the audit client's year-end falls in December or March (*BUSY*) and whether the company is located in London (*LOND*). The latter two variables are expected to be positively related to audit fees since companies audited during the 'busy' period may be charged higher fees due to the higher opportunity cost of audit resources (e.g. McMeeking et al., 2006) while companies located in London are expected to pay higher audit fees reflecting cost of living differentials (Chaney et al., 2004; Clatworthy and Peel, 2007).

Other than in respect of corporate size and complexity, the literature on the choice of variables in the auditor selection model is less developed and prior studies are usually based on including a subgroup of variables from the audit fee equation in the selection model (Chaney et al., 2004, 2005; Hamilton et al., 2005, but cf. Ireland and Lennox, 2002). If one assumes that firms choose auditor type by comparing their predicted costs (fees), the choice of auditor type depends on all the factors affecting the fees charged by either type of auditor. We therefore included all variables from the fee equations in the auditor choice model. When all or a subset of the regressors from the fees equation is used, identification relies on the non-linearity of the selection term but this non-linearity may not be sufficient to produce convincing estimates. It is important therefore to include an identification variable that is significantly associated with auditor choice (in the probit model), but not with audit fees (in the fees equation). Such variables are extremely hard to obtain in practice (see, e.g. Puhani, 2000). We considered several plausible instruments[12] and found only one – the change in the absolute value of total assets (*CHTA*) between the current and preceding year – which was statistically significant (with the expected sign) in the probit selection model, but statistically insignificant when included in the OLS audit fee models.[13] Furthermore it is not formally a 'weak' instrument since it has an $F$-statistic of 11.21 for the null that it is insignificant in the regression of auditor type ($D$) on all the regressors. This exceeds the critical value of 8.96 for the validity of a single instrument given by Stock et al. (2002) and the informal value of 10 that is widely used and advocated by Stock and Watson (2003: 350).

The motivation for *CHTA* being included in the selection model is that companies involved in large investments/acquisition or divestments/sale of assets, may require the expertise of a Big Four auditor due to the additional complexity of the audit. In addition, Keasey and Watson (1991) note that the absolute change in firm size (total assets) may, from an agency perspective, act as a proxy for contractual changes at the firm level, which

[11] Because company records on the database we use (FAME) only indicate whether or not either of these events occurred, we are unable to refine *PBAL* or *CONLIAB* to take account of the types of events or the nature of liabilities. Hence, we assess their average impact. We were also unable to ascertain the nature of the qualification and hence through *QUALIF*, we again measure the average impact of a qualified audit report.

[12] These included changes in sales, change in equity, change in total assets and various transformations of these variables.

[13] The $t$-values for the *CHTA* coefficients when included in models 1, 2 and 3 in Table 3 below were, respectively: 0.05, 1.32 and 0.18.

**Figure 5**
**Empirical model of auditor selection**

$$BIG4 = \delta_0 + \delta_1 \ln SAL + \delta_2 \ln TA + \delta_3 SQSUBS + \delta_4 EXPSAL + \delta_5 QUALIF + \delta_6 PBAL + \delta_7 CONLIAB \quad (8)$$

$$+ \delta_8 EXITEM + \delta_9 RTA + \delta_{10} TLTA + \delta_{11} LOND + \delta_{12} BUSY + \delta_{13} CHTA + \varepsilon$$

could prompt a change in the demand for auditing services. Hence, large auditors may be associated with reducing agency costs (e.g. Ireland and Lennox, 2002) in companies with large asset variations. Although it has desirable theoretical qualities, it is also employed for pragmatic reasons, since it formally fulfils its main purpose of properly identifying the audit fee equations. Our empirical model of auditor selection is shown in Figure 5.

Following previous studies (e.g. Chaney et al., 2004; Hamilton et al., 2005), we expect the variables reflecting auditee size (ln*SAL* and ln*TA*) and complexity (*SQSUBS* and *EXPSAL*) to be positively associated with the choice of a Big Four auditor in the probit model, in consequence of their hypothesised capacity to provide more efficient audits and to reduce agency costs (ibid.). In line with prior research (e.g. Ireland and Lennox, 2002; Chaney et al., 2004, 2005), we also expect our auditee risk variables (*QUALIF*, *PBAL*, *CONLIAB*, *EXITEM*, *RTA* and *TLTA*) to be positively associated with the selection of a Big Four auditor.[14] As noted by Hamilton et al. (2005: 9), 'The greater the client's risk, the higher the propensity for the impairment of agency relationships. To mitigate the associated agency costs, higher quality auditors, surrogated by big 4, are more likely to be selected to signal the credibility of reporting.' Furthermore, Datar et al. (1991) predict, and Copley and Douthett (2002) find, a positive relationship between auditee risk and the appointment of a higher quality auditor.

For the final two variables (*LOND* and *BUSY*), we have no strong prior expectations about their influence on auditor choice, although the univariate results of extant studies (as in the current study) have consistently reported (for both private and quoted audit clients) that a significantly higher proportion of Big Four auditors conduct their audits during the busy period, while a significantly higher proportion of non-Big Four auditors are appointed to companies located in London (e.g. Ireland and Lennox, 2002; Chaney et al., 2004).

### *3.2. Data*

The source of our data is the Bureau Van Dijk FAME DVD-ROM UK database. Financial data (annual accounts) and non-financial data (e.g. company location, auditor and audit qualification) are available as individual records for each company on the database. Companies were included if they met the following criteria: their primary activities (according to FAME primary SIC codes) were outside the financial sector; they were private limited companies; they were 'live' companies (i.e. had not ceased trading, failed or entered into voluntary liquidation); their audited accounts were available on FAME; they had full data available, including total assets and sales (minimum £1,000), audit fee (minimum £100), and a disclosed profit/loss figure. In order to avoid the potential confounding influences of including both holding companies and their subsidiaries in the regression model (e.g. Ezzamel et al., 1996; Peel and Roberts, 2003), our sample only includes independent companies (i.e. those not held as a subsidiary of another company). In line with previous studies (e.g. Firth, 1997), financial companies were excluded due to the different composition of their financial statements and only live companies were selected to avoid the confounding influence of including non-live auditees. In addition, and in line with previous research, 11 companies with joint auditors (none of which were Big Four auditors) were excluded from the analysis to comply with the binary nature of the probit model. Following these restrictions, we obtained the necessary data for a sample of 36,674 private companies from FAME for the latest financial statements available (predominantly for the calendar year 2003).[15] It is very important to note (since it has significant effects on both sample size and data accuracy) that the FAME default setting for downloading data is £000s, with data being rounded to the nearest £1,000; for example an audit fee of £1,550 would be rounded to £2,000 and one of £400 to zero (i.e.

[14] As noted by an anonymous reviewer, it is possible that if large auditors do not favour small companies due to lower return relative to the risk involved, small firms will have to pay more to retain a large auditor. This does not, however, preclude small companies paying such a premium if they believe the quality of audit to be higher for the reasons discussed above.

[15] A possible limitation of our work is that we employ a cross-sectional, rather than a panel-based sample. Prior research by Chaney et al. (2004) suggests that private company audit fee models do not vary much over time. Their sample covers a five-year period and there is little variation in their annual estimates; however, cf. McMeeking et al. (2006), who do find inter-temporal variation for listed companies.

**Table 2**
**Descriptive statistics**

| | *Big Four clients (n = 3,038)* | | | *Non-Big Four clients (n = 33,636)* | | | *Total sample (n = 36,674)* | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| *Variable* | *Mean* | *Std. dev.* | *Median* | *Mean* | *Std. dev.* | *Median* | *Mean* | *Std. dev.* | *Median* | *Sig.* |
| *AFEE* (£000) | 29.05 | 80.47 | 13.00 | 5.88 | 13.12 | 2.75 | 7.80 | 27.11 | 3.00 | ‡§ |
| ln*AFEE* | 9.44 | 1.25 | 9.47 | 7.91 | 1.22 | 7.92 | 8.04 | 1.29 | 8.01 | ‡§ |
| *SAL* (£m) | 39.41 | 150.62 | 8.14 | 5.13 | 21.46 | 0.84 | 7.97 | 48.89 | 1.02 | ‡§ |
| *TA* (£m) | 35.62 | 159.41 | 6.08 | 3.17 | 14.49 | 0.48 | 5.86 | 48.75 | 0.59 | ‡§ |
| *SUBS* | 3.46 | 8.37 | 1.00 | 0.61 | 2.46 | 0.00 | 0.84 | 3.46 | 0.00 | ‡§ |
| *EXPSAL* | 0.08 | 0.20 | 0.00 | 0.02 | 0.12 | 0.00 | 0.03 | 0.13 | 0.00 | ‡§ |
| *QUALIF* | 0.04 | 0.19 | 0.00 | 0.03 | 0.17 | 0.00 | 0.03 | 0.18 | 0.00 | |
| *PBAL* | 0.12 | 0.33 | 0.00 | 0.04 | 0.19 | 0.00 | 0.04 | 0.21 | 0.00 | ψ |
| *CONLIAB* | 0.27 | 0.45 | 0.00 | 0.10 | 0.30 | 0.00 | 0.11 | 0.31 | 0.00 | ψ |
| *EXITEM* | 0.57 | 0.50 | 1.00 | 0.32 | 0.47 | 0.00 | 0.34 | 0.47 | 0.00 | ψ |
| *RTA* | 0.01 | 0.38 | 0.03 | 0.23 | 0.78 | 0.07 | 0.21 | 0.76 | 0.07 | ‡§ |
| *TLTA* | 0.84 | 1.51 | 0.73 | 0.77 | 1.15 | 0.66 | 0.78 | 1.18 | 0.67 | ‡§ |
| *LOND* | 0.21 | 0.41 | 0.00 | 0.34 | 0.47 | 0.00 | 0.33 | 0.47 | 0.00 | ψ |
| *BUSY* | 0.56 | 0.50 | 1.00 | 0.44 | 0.50 | 0.00 | 0.45 | 0.50 | 0.00 | ψ |
| *CHTA* (£m) | 5.26 | 33.82 | 0.61 | 0.49 | 2.71 | 0.05 | 0.88 | 10.16 | 0.06 | ‡§ |

*Notes:*
Variable definitions are provided in Table 1.
‡ and § indicate means and distributions are significantly different between Big Four and non-Big Four clients at the 0.01 level in *t*-tests and Mann-Whitney tests respectively.
ψ indicates significant difference between Big Four and non-Big Four clients at the 0.01 level in a chi-squared test.

a missing value). Data can, however, be downloaded (as in the current study) in £ and hence neither data accuracy nor observations are lost using this option. The sampling consequences of this are not trivial, since downloading in £ captures a large number of smaller firms. For instance, Chaney et al. (2004), whose sample excludes many small companies due to the imprecision associated with downloading in £000, report Big Four concentration of 50% compared to 8% in our sample.

Descriptive statistics are presented in Table 2. The average audit fee (*AFEE*) for the whole sample ($n = 36,674$) amounted to £7.80k, with companies having mean sales (*SAL*) and total assets (*TA*) of £7.97m and £5.86m respectively. Sales range from a minimum of £1k to a maximum of £4,979m and total assets from £1k to £5,234m. Table 2 also shows that, other than audit qualifications (*QUALIF*), all variables differ significantly between the Big Four ($n = 3,038$) and the non-Big Four ($n = 33,636$) sub-samples. Consistent with prior expectations, Big Four clients are significantly larger (as measured by both *SAL* and *TA*), have more subsidiaries (*SUBS*), have a higher proportion of foreign to total sales (*EXPSAL*) and report more post balance sheet events (*PBAL*), contingent liabilities (*CONLIAB*) and exceptional items (*EXITEM*). In addition, Big Four clients are less profitable (*RTA*) more highly geared (*TLTA*), less likely to be located in London (*LOND*), more likely to be audited during the busy period (*BUSY*), with a significantly higher absolute change in the value of total assets (*CHTA*). Due to the large number of small auditees represented in the non-Big Four sample, the differences in size between Big Four (average sales and total assets of £39.41m and £35.62m) and non-Big Four auditees (average sales and total assets of £5.13m and £3.17m) are substantial.

## 4. Empirical results

We commence our analysis with standard single-stage OLS regression under the assumption of no selection bias. We then report our comparative analysis employing the two-step Heckman procedure, together with associated robustness tests. Finally, we present the results of the matching procedures.

### *4.1. Single stage results*

Model 1 in Table 3 shows the OLS estimates for the standard pooled audit fee specification, which is employed in many previous studies. All explanatory variables take their expected signs and other than the busy period variable (*BUSY*), which is statistically significant at the 0.10 level ($p = 0.079$), all are highly significant ($p < 0.0001$ in all cases).

**Table 3**
**Regression results**

| | OLS single stage models | | | MLE two-step models | | | Heckman two-step models | | |
|---|---|---|---|---|---|---|---|---|---|
| | *Model 1 (Pooled fee)* | *Model 2 (Big Four fee)* | *Model 3 (Non-Big Four fee)* | *Model 4a (Probit choice)* | *Model 5a (Big Four fee)* | *Model 6a (Non-Big Four fee)* | *Model 4b (Probit choice)* | *Model 5b (Big Four fee)* | *Model 6b (Non-Big Four fee)* |
| ln*SAL* | 0.284 | 0.285 | 0.286 | –0.004 | 0.285 | 0.286 | –0.005 | 0.285 | 0.285 |
| | (87.31)** | (25.08)** | (83.84)** | (0.48) | (30.93)** | (96.44)** | (0.50) | (29.51)** | (92.71)** |
| ln*TA* | 0.122 | 0.120 | 0.120 | 0.246 | 0.149 | 0.113 | 0.246 | 0.211 | 0.101 |
| | (36.28)** | (10.30)** | (33.54)** | (23.25)** | (8.00)** | (34.86)** | (23.20)** | (5.41)** | (25.46)** |
| *SQSUBS* | 0.258 | 0.201 | 0.281 | 0.087 | 0.209 | 0.268 | 0.088 | 0.227 | 0.247 |
| | (44.70)** | (19.07)** | (40.10)** | (6.98)** | (19.81)** | (46.05)** | (7.03)** | (15.28)** | (34.78)** |
| *EXPSAL* | 0.367 | 0.627 | 0.293 | 0.607 | 0.688 | 0.253 | 0.602 | 0.822 | 0.191 |
| | (13.15)** | (10.91)** | (9.41)** | (9.84)** | (11.34)** | (8.98)** | (9.77)** | (8.36)** | (6.17)** |
| *QUALIF* | 0.115 | 0.141 | 0.111 | –0.146 | 0.126 | 0.118 | –0.146 | 0.094 | 0.128 |
| | (6.18)** | (2.56)* | (5.58)** | (2.61)** | (2.30)* | (6.06)** | (2.61)** | (1.57) | (6.36)** |
| *PBAL* | 0.119 | 0.179 | 0.098 | 0.169 | 0.196 | 0.084 | 0.169 | 0.233 | 0.063 |
| | (7.74)** | (5.65)** | (5.62)** | (4.23)** | (5.85)** | (4.69)** | (4.23)** | (5.62)** | (3.34)** |
| *CONLIAB* | 0.095 | 0.064 | 0.099 | 0.014 | 0.064 | 0.096 | 0.013 | 0.066 | 0.091 |
| | (8.93)** | (2.64)** | (8.47)** | (0.48) | (2.57)* | (7.68)** | (0.45) | (2.45)* | (7.09)** |
| *EXITEM* | 0.131 | 0.126 | 0.130 | –0.081 | 0.118 | 0.133 | –0.080 | 0.100 | 0.137 |
| | (17.04)** | (5.57)** | (15.88)** | (3.38)** | (5.06)** | (16.14)** | (3.35)** | (3.78)** | (15.99)** |
| *TLTA* | 0.026 | –0.009 | 0.029 | 0.076 | –0.004 | 0.025 | 0.076 | 0.007 | 0.019 |
| | (6.43)** | (1.31) | (6.24)** | (9.07)** | (0.48) | (8.04)** | (8.98)** | (0.59) | (5.38)** |
| *RTA* | –0.033 | –0.111 | –0.031 | –0.231 | –0.139 | –0.031 | –0.230 | –0.199 | –0.032 |
| | (7.14)** | (3.28)** | (6.65)** | (8.12)** | (3.87)** | (6.82)** | (8.10)** | (4.02)** | (6.54)** |
| *LOND* | 0.208 | 0.338 | 0.200 | –0.306 | 0.306 | 0.210 | –0.305 | 0.237 | 0.226 |
| | (29.69)** | (12.38)** | (27.63)** | (12.15)** | (9.93)** | (29.10)** | (12.11)** | (4.72)** | (27.93)** |
| *BUSY* | 0.011 | 0.010 | 0.010 | 0.153 | 0.026 | 0.004 | 0.152 | 0.0601 | –0.005 |
| | (1.76) | (0.52) | (1.51) | (7.12)** | (1.18) | (0.63) | (7.10)** | (2.01)* | (0.65) |

**Table 3**
**Regression results** *(continued)*

| | OLS single stage models | | | MLE two-step models | | | Heckman two-step models | | |
|---|---|---|---|---|---|---|---|---|---|
| | *Model 1 (Pooled fee)* | *Model 2 (Big Four fee)* | *Model 3 (Non-Big Four fee)* | *Model 4a (Probit choice)* | *Model 5a (Big Four fee)* | *Model 6a (Non-Big Four fee)* | *Model 4b (Probit choice)* | *Model 5b (Big Four fee)* | *Model 6b (Non-Big Four fee)* |
| *CONSTANT* | 2.299 (88.49)** | 2.638 (23.07)** | 2.302 (83.80)** | –4.886 (46.58)** | 1.967 (5.14)** | 2.378 (87.60)** | –4.870 (46.62)** | 0.511 (0.58) | 2.499 (69.11)** |
| *BIG4* | 0.270 (22.96)** | | | | | | | | |
| *CHTA* | | | | 0.007 (4.17)** | | | 0.007 (4.27)** | | |
| *IMR* (λ) | | | | | 0.142 (1.80) | –0.199 (9.56)** | | 0.446 (2.42)** | –0.509 (8.39)** |
| *N* | 36,674 | 3,038 | 33,636 | 36,674 | 3,038 | 33,636 | 36,674 | 3,038 | 33,636 |
| Adj. $R^2$ | 0.78 | 0.80 | 0.75 | – | – | – | – | 0.80 | 0.75 |

*Notes:* This table reports regression estimates where the dependent variable is lnAFEE for all models, except Models 4a and 4b where the binary dependent variable is coded 1 if the firm is audited by a Big Four firm, 0 otherwise. Variable definitions are provided in Table 1. Absolute *t*-statistics are reported in parentheses, except for Models 4b–6b, where *z*-statistics are in parentheses; *t*-statistics for OLS models use robust standard errors and corrected standard errors for the Heckman two-step models. * and ** indicate statistical significance at 0.05 and 0.01 levels respectively.

**Figure 6**
**OB decomposition**

$$\overline{\ln F_{BIG4}} - \overline{\ln F_{NON}} = \sum_{k=1}^{K} a_k(\overline{X_{BIG4k}} - \overline{X_{NONk}}) + b + \sum_{k=1}^{K}(b_k - a_k)\overline{X_{BIG4k}} \quad (9)$$

Actual difference = Explained by characteristics + Big Four premium

$$9.4364 - 7.9099 = 9.1809 - 7.9099 + 9.4364 - 9.1809 \quad (10)$$

$$\underset{}{1.5265} = \underset{(263.6)}{1.2709} + \underset{(21.87)}{0.2556}$$

In particular, we note that the *BIG4* coefficient (0.270) implies that, on average, the audit fees of a non-Big Four auditee would increase by 31% if it were to employ a Big Four auditor.[16] Also noteworthy is that the model explains a relatively high proportion ($R^2$ of 78%) of the variation in the audit fees of UK private companies, comparing favourably with that (57%) reported by Chaney et al. (2004) for their sample of UK private firms.

Models 2 and 3 in Table 3 report OLS estimates for separate audit fee equations for the Big Four and non-Big Four auditee samples. In Model 1, the Big Four equation only differs by a constant from the non-Big Four equation. In common with Chaney et al. (2004) a joint F-test rejected the null hypothesis ($F = 13.43$; $p = 0.000$) that the coefficients in the Models 2 and 3 were the same, implying that the fee-setting process differs between the two auditor types. The main focus of our empirical analysis is therefore Models 2 and 3 in Table 3 (i.e. those which allow the slope coefficients of the explanatory variables to differ for the Big Four and non-Big Four models).

Table 3 shows that for the non-Big Four specification (Model 3) all explanatory variables exhibit their expected signs and, other than for *BUSY*, which loses statistical significance ($p = 0.131$), all variables are highly significant ($p < 0.0001$ in all cases). For the Big Four specification (Model 3), in addition to *BUSY*, the sign on the gearing coefficient (*TLTA*) is negative, but statistically insignificant – a finding in common with Chaney et al. (2004) for their Big Four equation; furthermore, the intercept in Model 2 is larger than in Model 3 – a result also reported by Chaney et al. (2004) and attributed to Big Four auditors recovering higher expenditure on training and facilities.

To examine the premium in more detail, we conduct the Oaxaca-Blinder decomposition (discussed above) on our estimates for Models 2 and 3. The OB decomposition is based on measuring the premium (using the characteristics of the average Big Four auditee) as in Figure 6.

Greene (2003: 54) provides the formulae for the estimated standard errors of each term in the decomposition and we report the $t$-values based on this method in parentheses under the estimates. There is a large and significant ($p = 0.000$) difference in the means of the audit fees paid by companies audited by Big Four and non-Big Four auditors (1.5265) using the parameters from Models 2 and 3. Most of this is accounted for by differences in their respective client characteristics (1.2709 or 83%). However, there is, on average, a significant ($p = 0.000$) Big Four premium of 0.2556 (29.1%), which is close to that (31.0%) estimated in the pooled OLS equation (Model 1) and in line with findings in prior research (e.g. Moizer, 1997). On average, Big Four auditees paid audit fees of £12,537 ($e^{9.4364}$), but would have paid £9,710 if they were charged according to the non-Big Four parameters (Model 3) – a reduction of 23%.[17] Hence, the Oaxaca-Blinder results, based on Models 2 and 3 are consistent with the presence of a Big Four audit premium. The next section presents our two-stage results where we analyse the extent to which these findings are affected by selection bias.

### *4.2. Heckman two-step regressions*

Table 3 reports the two-step results with maximum likelihood estimates (MLE) and standard Heckman two-step estimates. Models 4a and 4b show the probit selection model estimates for the choice of a Big Four auditor, while Models 5a (5b) and 6a (6b) report the MLE (standard Heckman) audit fee regression estimates for the Big Four and non-Big Four auditees, including the additional parameter $\lambda$ (for the IMR estimated from the coefficients in Model 4) to control for selection bias. The MLE and standard Heckman two-step probit selection models (4a and 4b) are very similar, and in both models, all explanatory variables other

[16] We use the standard transformation $e^x - 1$ (where $x$ = the coefficient or mean log difference) to compute percentages.

[17] The alternative decomposition using the characteristics of non-Big Four clients also implied a statistically significant premium (at $p = 0.000$) of 31%.

**Figure 7**
**Predicted fees**

$$\overline{\ln F_{BIG4}} = a + b + \sum_{k=1}^{K} b_k \overline{X_{k\,BIG4}} + g_{SBIG} \overline{\lambda_{BIG4,BIG4}} \quad \text{For Big Four clients (actual)} \qquad (11)$$

$$\overline{\ln F_{NON}} = a + \sum_{k=1}^{K} a_k \overline{X_{k\,BIG4}} + g_{SNON} \overline{\lambda_{BIG4,BIG4}} \quad \text{For Big Four clients (counterfactual)} \qquad (12)$$

than *CONLIAB* and ln*SAL* are significantly associated with auditor choice at the 0.01 level. In particular, the coefficient on the identifying variable (*CHTA*) exhibits its expected sign and is highly statistically significant ($p = 0.000$).[18] Also consistent with expectations and prior research, Models 4a and 4b show that larger (ln*TA*), more complex (*SQSUBS*; *EXPSAL*) and riskier (*RTA*; *TLTA*) companies are more likely to appoint a Big Four auditor. Companies receiving audit qualifications (*QUALIF*) are more likely to employ a non-Big Four auditor, in contrast to companies reporting a post-balance sheet event (*PBAL*) and auditees based in London, which are less likely to select a Big Four auditor, although likely to do so if their year ends fall in the busy period (*BUSY*).[19]

The audit fee equations (Models 5a, 6a, 5b and 6b) contain the same pattern of significance levels as the single stage estimates in Models 2 and 3. The MLE estimates in Table 3 show that the λ coefficient is negative but highly significant ($p = 0.000$) for the non-Big Four equation (Model 6a), but positive and only significant at the 0.10 level ($p = 0.071$) in the Big Four equation (Model 5a). The positive MLE estimate of 0.142 (Model 5a) for the covariance $\sigma_{SBIG4}$ and the negative estimate of –0.199 (Model 6a) for $\sigma_{SNON}$ imply that an increase in the value for the unobservable error in the auditor selection equation ($\varepsilon_{SEL}$) is associated with an increase in the value of the unobservable component of Big Four fees ($\varepsilon_{BIG4}$) and a decrease in the value of unobservable component of non-Big Four fees ($\varepsilon_{NON}$), although the former estimate is insignificant at the 0.05 level. These results imply that the effect of unobservable auditee characteristics is to cause private companies to choose the most expensive auditor and directly contradict the results of Chaney et al. (2004). The results also imply non-Big Four auditees value each type of auditor differently from Big Four auditees: not only are they willing to pay more for non-Big Four auditors; they also place a lower value on the services of a Big Four auditor.[20]

The selection estimates of the Big Four premium are also dependent on the estimator used. As the results in Models 5b and 6b show, the standard Heckman two-step approach amplifies the MLE estimates. Although the λ coefficients have the same signs as their MLE counterparts, and are both significant at the 0.05 level in both equations, they are implausibly large (in absolute terms) at 0.446 for the Big Four and –0.509 for the non-Big Four auditees; thus the λ coefficient in the Big Four equation more than doubles (compared to the MLE). Further indications of model instability are provided by the insignificance of the intercept and *QUALIF* in the Big Four Heckman model (5b).

It has been acknowledged in prior research (Stolzenberg and Relles, 1997) that interpreting the magnitude of the λ coefficient is difficult due to the abstract nature of the variable itself. Extending the logic of our earlier decomposition to the Heckman results is informative in this context since it allows an assessment of the *effects* of the coefficient. To calculate the impact of selection bias on the Big Four premium by decomposing the observable and unobservable effects, we concentrate on the Big Four premium measured at the sample means of the Big Four auditees, i.e. the average effect of the treatment on the treated (*ATT*). Predicted fees paid by a Big Four client at the sample means are shown in Figure 7.

[18] The statistical insignificance of ln*SAL* in the auditor choice equation is not related to collinearity with *CHTA*. When *CHTA* was removed from Model 4, ln*SAL* remained statistically insignificant. In addition, when ln*SAL* was removed from Model 4, *CHTA* remained positive and statistically significant.

[19] The Wald statistic of 2668.40 ($p < 0.0001$) for Model 4b indicates the selection equation is well determined; the McFadden's $R^2$ is 0.204 and the model correctly classifies (cut-off point of 0.083 – representing the prior probability of selection into the Big Four) 77.52% and 70.81% of the Big Four and non-Big Four auditees respectively.

[20] Although *a priori* these findings may appear counterintuitive and may relate, as discussed below, to the lack of robustness of Heckman procedures in audit fee studies, they are not entirely implausible. As discussed earlier, there are numerous explanations (e.g. audit quality effects) for firms paying higher fees for Big Four audits; similarly, survey-based research by Marriott et al. (2007) finds that very small UK companies prefer non-Big Four auditors due to the more personal services and stronger relationships offered by smaller auditors. A further possibility is that the potential financial gains arising from switching auditor may not justify the associated costs.

**Figure 8**
**Estimates for Models 5a and 6a**

$$\overline{\ln F_{BIG4}} = a + b + \sum_{k=1}^{K} b_k \overline{X_{k\,BIG4}} + g_{SBIG} \overline{\lambda_{BIG4,BIG4}} \qquad \text{For Big Four clients (actual)} \qquad (13)$$

$$9.4364 = 9.2268 + 0.2096$$

And:

$$\overline{\ln F_{NON}} = a + \sum_{k=1}^{K} a_k \overline{X_{k\,BIG4}} + g_{SNON} \overline{\lambda_{BIG4,BIG4}} \qquad \text{For Big Four clients (counterfactual)} \qquad (14)$$

$$8.8230 = 9.1174 - 0.2944$$

**Table 4**
**Effects of changes in specification on MLE and standard Heckman two-step models**

| | *Specification 1: full fee equation; full selection equation* | | *Specification 2: full fee equation; selection equation excludes CHTA* | |
|---|---|---|---|---|
| | *MLE* | *Heckman* | *MLE* | *Heckman* |
| Big Four λ coefficient | 0.142 | 0.446 | 0.166 | 1.079 |
| | (1.80) | (2.42)* | (2.31)* | (3.74)** |
| Non-Big Four λ coefficient | –0.199 | –0.509 | –0.204 | –0.595 |
| | (9.56)** | (8.39)** | (10.00)** | (9.22)** |
| Big Four conditional mean$^{\Psi}$ | 9.436 | 9.436 | 9.436 | 9.436 |
| Non-Big Four conditional mean | 8.823 | 8.263 | 8.812 | 8.107 |
| Difference (Big Four premium as *ATT*) | 0.613 | 1.1729 | 0.624 | 1.329 |
| Big Four unconditional mean | 9.227 | 8.776 | 9.189 | 7.834 |
| Non-Big Four unconditional mean | 9.117 | 9.018 | 9.115 | 8.989 |
| Difference (Big Four premium as *ATE*) | 0.109 | –0.242 | 0.074 | –1.155 |
| | (0.93) | (0.88) | (0.69) | (2.69)** |
| Big Four selection effect | 0.210 | 0.660 | 0.246 | 1.602 |
| Non-Big Four selection effect | –0.294 | –0.755 | –0.303 | –0.883 |
| Difference in selection effect | 0.504 | 1.415 | 0.549 | 2.485 |
| | (4.20)** | (4.92)** | (4.95)** | (5.66)** |
| Big Four $R^2$ (dep. var. = λ)$^{\S}$ | 0.986 | 0.986 | 0.992 | 0.992 |
| Non-Big Four $R^2$ (dep. var. = λ)$^{\S}$ | 0.854 | 0.858 | 0.868 | 0.874 |

*Notes:*
The table reports maximum likelihood (MLE) and standard Heckman estimates for three specifications of the two-step correction model. Specification 1 corresponds to the MLE and Heckman models reported in Table 3. Absolute *t*-values are reported in parentheses.
**, * represent statistical significance at the 0.01 and 0.05 level, respectively.
$^{\Psi}$ The conditional mean is the predicted mean of audit fees conditional on the auditee choosing a Big Four or non-Big Four auditor, allowing for auditees' unobservable characteristics. The unconditional mean is the predicted mean of audit fees excluding the selection effects. The difference between the conditional and unconditional means is the selection effect.
$^{\S}$ reports the $R^2$ for a regression of the selection term on the remaining variables in the audit fee (second stage) equation.

**Figure 9**
**Relationship between treatment effects**

$$\overline{\ln F_{BIG4,BIG4}} - \overline{\ln F_{NON,BIG4}} = b + \Sigma_k (b_k - a_k)\overline{X_{kBIG4}} + g_{SBIG}\overline{\lambda_{BIG4,BIG4}} - g_{SNON}\overline{\lambda_{BIG4,BIG4}} \quad (15)$$

$(ATT) = (ATE) + \text{Estimate}[E(\varepsilon_{BIG4} \mid D = 1) - E(\varepsilon_{NON} \mid D = 1)]$

| Difference in conditional means | = | Difference in unconditional means | + | Difference in unobservable effect |
|---|---|---|---|---|
| 9.4364 – 8.8230 | = | 9.2268 – 9.1174 | + | 0.2096 + 0.2944 |
| 0.6134 | = | 0.1094 | + | 0.504 |

The counterfactual equation shows the predicted audit fees for a typical Big Four client which paid audit fees according to the non-Big Four model. Since the same regressor means are used to compute predicted audit fees, we have removed any potential differences due to the different characteristics (the explained differences) of the Big Four and non-Big Four auditees, with any remaining difference amounting to the Big Four premium (the unexplained differences). The two components of predicted audit fees estimate the separate effects of the observable regressors and the unobservables. The decomposition of the counterfactual audit fees ($\overline{\ln F_{NON}}$) comprises the predicted fees paid to a non-Big Four auditor by any firm with the same mean observable characteristics

$$(a + \sum_{k=1}^{K} a_k \overline{X_{kBIG4}})$$

plus the selection effect ($g_{SNON}\overline{\lambda_{BIG4,BIG4}}$) showing the predicted effect of unobservable characteristics. The first term represents the unconditional mean showing the predicted audit fees if the clients chose Big Four and non-Big Four auditors at random.[21] The predicted fees ($\overline{\ln F}$) incorporate the selection terms and are therefore referred to as the conditional means.[22] Hence, the predicted fee equations have the form: conditional mean = unconditional mean + selection effect. The estimates for Models 5a and 6a in Table 3 are shown in Figure 8.

These estimates, along with those for non-Big Four firms, are tabulated in Table 4. Thus the typical Big Four auditee actually paid fees in natural log form of 9.4364 (£12,537). By contrast it would have paid predicted fees as a non-Big Four auditee of 8.8230 in natural log form (£6,789), implying a Big Four premium of 0.6134 or 85%, which is much larger than that found in prior research. Our MLE results suggest that on average, Big Four auditees would have paid 9.2268 (£10,166) for the services of a Big Four auditor and 9.1174 (£9,113) for a non-Big Four auditor if their unobservable characteristics were ignored. However, in consequence of Big Four auditees' unobservable characteristics, an additional 0.2096 is paid for Big Four audit fees, and 0.2944 less for the services of a non-Big Four auditor.

The average treatment effect on the treated (*ATT*) is the difference in the conditional means of the audit fees paid by Big Four and non-Big Four auditees and represents the difference in fees only available to Big Four auditees. By contrast, the average treatment effect (*ATE*) shows the difference in fees available to any auditee. The relationship between the treatment effects is shown in Figure 9 (see Heckman et al., 2001).

Using the MLE parameters, the Big Four premium paid by Big Four auditees or *ATT* is 0.6134 (85%). The typical Big Four auditee paid 0.1094 (12%) more in fees based on their observable characteristics. This *ATE* (12%) is assumed to be freely substitutable, in that any non-Big Four auditee with the relevant characteristics who switched to a large auditor would incur this premium and vice versa. However the peculiar unobservable characteristics of Big Four auditees (as reflected in their value of $\lambda$) mean that they would pay an additional 0.504 in natural log terms for the services of a Big Four auditor, whereas the unobserved characteristics of non-Big Four auditees means they would be unwilling to pay this premium. Since the selection effects are individually significant, they should be included in the model. Although the *ATE* is not economically insubstantial at 12%, it is not significantly different from zero ($t = 0.93$; $p = 0.35$). By contrast the large selection effect is highly significant ($t = 4.20$; $p = 0.00$). According to our MLE results, therefore, firms with similar observable characteristics would pay higher fees if they used

[21] Note that with random selection, there would be no selection effect.

[22] They are conditional in the sense that they are estimates of the expected audit fees conditional on the firm employing either a Big Four or a non-Big Four auditor.

a Big Four auditor, but the difference is statistically insignificant. However, auditees differ greatly in their unobserved characteristics and these differences largely generate the Big Four premium. We emphasise that our results contrast sharply with Chaney et al. (2004), who report that unobservable factors make it *cheaper* for Big Four auditees to opt for Big Four auditors, rather than non-Big Four ones.

The Heckman two-step results in Table 4 (Specification 1) show that the typical Big Four auditee actually paid log fees of 9.436 (£12,537). By contrast it is predicted to have incurred fees as a non-Big Four auditee of 8.263 (£3,880) estimating the Big Four premium at an inconceivable 1.1729 or 223% (compared to 85% for the MLE estimates discussed above). Big Four auditees would have paid 8.776 (£6,478) for the services of a Big Four auditor and 9.018 (£8,250) for a non-Big Four auditor, if their unobservable characteristics were ignored. However, as a result of Big Four auditees' unobservable characteristics, an additional 0.660 (Big Four selection effect) is paid for the services of a Big Four auditor, and 0.755 less for the services of a non-Big Four auditor (non-Big Four selection effect). The effect of unobservables is to increase the Big Four fees (in £) by 93% ($e^{0.660}$=1.93). These estimates imply that Big Four auditees choose the cheaper auditor based on observed information, but the most expensive when unobservable characteristics are taken into account. Moreover, there are substantial differences between the MLE and Heckman estimates, which in itself, is indicative of model instability.

To assess the potential for parameter instability, Table 4 also examines the consequences of omitting the exclusion variable (*CHTA*) from the probit auditor selection equation. In this instance, changes in the MLE estimates are relatively minor, with the premium paid by the Big Four increasing moderately from 0.613 (85%) to 0.624 (87%). The difference in unconditional means (*ATE*) in the Heckman model, however, increases almost fivefold to –1.155 and is now statistically significant (at $p < 0.01$), although this is more than offset by the estimated impact of the unobservable selection difference (2.485). Comparing specifications therefore suggests the Heckman results are highly sensitive to model specification using the standard approach, although this is apparently less problematic when estimated using maximum likelihood.

The $R^2$ reported for both Big Four and non-Big Four equations in the bottom two rows of Table 4 for both models also testify to the high levels of multicollinearity when the $\lambda$ variable is regressed on the other variables in the audit fee equation. In Specification 1, these are 0.986 and 0.854 (i.e. variance inflation factors of 71.4 and 6.8) for the Big Four and non-Big Four equations respectively, suggesting that even *with* an identifying variable (*CHTA*) multicollinearity may pose problems for the Big Four estimation. To summarise the results from Table 4, the ML estimator appears most efficient whereas the estimates provided by the standard Heckman two-step procedure (the only method used for dealing with selection bias in the auditing literature to date) are potentially seriously unstable.

In Table 5, we examine the sensitivity of the Heckman and ML estimates of the selection term ($\lambda$) by changing the samples and variables used to estimate the selection models; we also report estimates of the premium using a single stage OLS pooled model for comparison, together with the associated $R^2$ and sample size ($n$) for the various models.

The sample in row 1 is based on all Big Four and all non-Big Four auditees and is reported for comparative purposes. Although it was not our aim to replicate the Chaney et al. (2004) study, as a further robustness test, row 2 reports estimates for the current sample, using the specification employed by Chaney et al. in their study of UK private companies (and noting the absence of a valid identifying variable in the selection equation).[23] The Chaney et al. (2004) model estimated in our sample has a substantially higher explanatory power ($R^2$ = 0.68 for the pooled model) than reported in their study (0.57), although this is still substantially lower than for the model specification (row 1) in this study (0.78)[24,25]. Of more importance, however, is that the Heckman two-step estimates of the $\lambda$ coefficients (row 2), are both highly significant ($p < 0.001$) and positive (1.850) and negative (–2.378) for the Big Four and non-Big Four models respectively. Similar findings are also evident from the MLE results. Hence, in contrast to Chaney et al. (2004), their models in our sample

[23] We are grateful to an anonymous reviewer for this suggestion. The variables are ln*TA*, *EXPSAL*, *LOND*, BUSY (all defined as in the current study), the ratio of exceptional and extraordinary items to total assets (*EXTA*), long-term debt to total assets, sales to total assets, the quick ratio, current assets to total assets, earnings before interest and taxes divided by total assets (*RTA*), *BUSY* × ln*TA* (*BTA*) and *RTA* × a loss indicator variable where unity = a company which made a loss in the prior period, zero otherwise. All variables were included in their audit fee model; but their selection model excluded *BUSY*, *BTA*, *LOND* and *EXTA*. There is a small fall in the sample size as a result of missing observations for the prior period loss indicator.

[24] Similar points also apply to the $R^2$ for separate Big Four and non-Big Four regressions.

[25] The lower explanatory power of the Chaney et al. (2004) models in their, relative to our, sample may partly relate to the larger sample (across all size ranges) employed in the current study and/or the rounding imprecision in their data noted earlier.

**Table 5**
**Effect of changes in sample and variables on OLS, Heckman two-step and MLE results**

| *Sample and variables* | *Pooled OLS* | | *Heckman two-step* | | | | *MLE* | |
|---|---|---|---|---|---|---|---|---|
| | | | *Big Four* | | *Non-Big Four* | | *Big Four* | *Non-Big Four* |
| | *BIG4* coefficient ($t$-value) | $R^2$ ($n$) | $\lambda$ coefficient ($t$-value) | $R^2$ ($n$) | $\lambda$ coefficient ($t$-value) | $R^2$ ($n$) | $\lambda$ coefficient ($z$-value) | $\lambda$ coefficient ($z$-value) |
| 1. Current sample; current variables | 0.270** (22.96) | 0.78 (36,674) | 0.446** (2.42) | 0.80 (3,038) | –0.509** (8.39) | 0.75 (33,636) | 0.142 (1.80) | –0.199** (9.56) |
| 2. Current sample; Chaney et al. (2004) variables | 0.386** (27.10) | 0.68 (36,607) | 1.850** (3.71) | 0.71 (3,026) | –2.378** (11.94) | 0.67 (33,581) | 0.277** (3.65) | –0.298** (19.67) |
| 3. Sample firms with total assets ≥ £10,000; current variables | 0.259** (21.99) | 0.76 (34,351) | 0.351* (2.38) | 0.80 (3,023) | –0.313** (4.84) | 0.73 (31,328) | 0.140 (1.86) | –0.193** (8.29) |
| 4. Thicker support region $^{\varphi}$; current variables | 0.280** (21.24) | 0.72 (33,007) | 0.075 (0.31) | 0.69 (2,188) | –0.524** (5.08) | 0.71 (30,189) | 0.020 (0.16) | –0.179** (6.78) |
| 5. Approximate matching sample$^{\Psi}$; current variables | 0.226** (18.91) | 0.69 (24,505) | 0.657 (1.81) | 0.76 (2,734) | –0.025 (0.24) | 0.66 (21,770) | 0.065 (0.74) | –0.290** (–7.56) |

*Notes:*
**, * represent statistical significance at the 0.01 and 0.05 level respectively; $t$-statistics for OLS models use robust standard errors and corrected standard errors for the Heckman two-step models.
All models labelled 'current variables' use the full model specification for the probit and audit fee equations in Table 3.
$^{\varphi}$ Thicker support region is represented by a sub-sample of firms where the top and bottom 5% of all sample firms are eliminated according to their probability of choosing a Big Four auditor.
$^{\Psi}$ Approximate matching sample is created by eliminating firms whose probability of selecting a Big Four auditor is greater (less) than the 95th (5th) percentile of the Big Four firms' probability of choosing a Big Four auditor.

produce the same signs on the λ coefficients for the Big Four and non-Big Four equations as those reported for the specifications employed in the current study (row 1) – although these coefficients are inconceivably large, being over four times larger than the respective coefficients of our Heckman two-step specifications. Furthermore, a contemporaneous working paper by Francis and Lennox (2008) reproduces the analysis of Chaney et al. (2004) using precisely the same variables and sample design (although for a later time period) confirms this instability. Inter alia, Francis and Lennox report an average positive λ coefficient (0.10) for the Chaney et al. Big Four auditor specification (as in the current study) and a positive one (0.21) for the non-Big Four specification. However, neither of the coefficients was statistically significant at conventional levels, suggesting no evidence of significant auditor selection bias. Finally, we note that using the Chaney et al. (2004) variables in our samples leads to a much higher OLS pooled estimate of the Big Four premium (47%) than that utilising the variables in this study (31%).

Since our auditee size cut-off point is low relative to previous studies, in row 3 we report parameter estimates for our previously reported models, but restricting the sample to firms with total assets exceeding £10,000.[26] The significance and signs of the λ coefficients are relatively stable – particularly the MLE parameters – although the $R^2$ of the pooled model declines (to 0.76) as does the estimated Big Four premium (to 29%).

In row 4 of Table 5, labelled the 'thicker support region', we exclude extreme observations (below the 5th and above the 95th percentiles) of the probability of selecting a Big Four auditor for the whole sample to provide estimates where there is more common support between the sub-samples. Row 5 excludes observations with a probability of selecting a Big Four auditor below the 5th, and above the 95th percentile for Big Four auditees and hence provides estimates where the Big Four and non-Big Four samples are more closely matched in terms of their selection probabilities. In row 4, the Big Four selection effect is now statistically insignificant for both the Heckman two-step and MLE parameters. In row 5, it is relatively large for the Heckman model but insignificant, as it is for the MLE model. These results underline the importance of comparable samples but at the possible cost of increasing multicollinearity. The non-linearity of the selection term is most noticeable for firms with high probabilities of selecting a Big Four auditor, so excluding these firms will tend to produce samples where a linear equation provides a more complete summary of the variation in λ. The converse applies for the non-Big Four firms and this is consistent with the Heckman two-step result in row 5, where large numbers of firms with high probabilities of choosing non-Big Four auditors are excluded.

In summary, unlike the standard single-stage pooled regression estimates of the Big Four premium, which Table 5 shows are always highly statistically significant in all models, this sensitivity analysis indicates that our original two-step results are not robust across all the sample and variable changes investigated. The sensitivity of the estimated selection effect to the omission of our identifying variable is consistent with extant econometric studies (above), and is of key importance to existing and future accounting studies which do not include a valid instrument when employing Heckman two-step procedures. The results in Table 5 also suggest that the positive selection effects for Big Four auditees may be driven by large companies and, with less certainty, the negative ones for non-Big Four auditees by small firms. There is evidence of serious instability in the coefficients, most noticeably varying by estimation technique. Nonetheless, the pattern of results gives no support to the hypothesis that Big Four auditees are choosing the lowest cost auditor after controlling for selection bias.

### 4.3. Matching results

Because of the sensitivity of the Heckman model demonstrated above, this section reports the results of the Big Four premium matching analysis. Under the conditional independence assumption, this produces estimates of the average treatment effect on the treated (*ATT*) not prone to the model specification and identification problems discussed above. We employ two matching methods: propensity score matching and a preprocess matching analysis combined with OLS regressions.

*Propensity score matching*

As discussed above, recent developments in the statistics and econometrics literature have suggested propensity score matching as an additional or alternative approach to two-step Heckman procedures. Since the seminal paper of Rosenbaum and Rubin (1983), propensity score matching has received considerable attention as a means of estimating causal treatment effects. In our analysis, Big Four auditees are matched to non-Big Four auditees on the basis of the predicted probability of employing a Big Four auditor – with the propensity scores (predicted probabilities) being derived from the probit selection equation (Model 4, Table

[26] We are grateful to the anonymous reviewers for suggesting this extension of our analysis. We also conducted this analysis for firms above and below the median and restricting the sample to firms with total assets over £1,000,000 and obtained further evidence of instability.

3), which includes all explanatory variables listed in Table 1.[27] Following matching, the average audit fees paid by the matched sub-samples are compared, to assess whether a significant premium is evident.[28]

The most popular propensity score matching method is nearest neighbour matching. In our study, this entails matching each Big Four auditee to the non-Big Four counterpart with the propensity score closest in value to that of the Big Four auditee. Nearest neighbour matching can be implemented with or without a 'calliper', where the calliper represents the maximum (absolute) difference between the propensity score of the nearest neighbour matched observations. A tighter calliper results in more closely matched observations but reduces the sample size, i.e. it only selects observations that can be matched within the minimum distance imposed by the calliper. When employing this method, researchers face a choice of whether to use replacement observations, i.e. permitting the use of non-Big Four (non-treated) auditees for matching with their Big Four counterparts more than once. This can be important in the nearest neighbour (without calliper) approach, since very large Big Four clients may have a limited number of counterparts in the non-Big Four sample; hence, excluding replacement can result in relatively large differences in propensity scores between the matched observations.

We therefore report in Table 6 (Panel A) results based on four different matching approaches to examine the robustness of our results and to illustrate the differences between the various methods,[29] each of which produces an equal number of matched Big Four and non-Big Four auditees. Panel A of Table 6 shows that the Big Four premium (the difference in the means of ln*AFEE* of the Big Four and non-Big Four sub-samples) is statistically significant at $p < 0.01$ under each type of matching, ranging from 0.2531 (28.8%) with a calliper of 0.001 (column 4) to 0.3082 (36.1%) under the nearest neighbour method with no replacement in column 3.[30] Moreover, as the statistics in column 5 demonstrate, even when Big Four and non-Big Four companies are very closely matched (with a maximum absolute difference in propensity scores of only 0.0001), the premium (0.2613 or 29.9%) remains robust and is within the range of premiums reported in prior research and in our earlier analysis based on OLS.

Although the results in column 5 demonstrate that the samples are very closely matched on their observed characteristics based on the composite propensity score, it is also important to examine how closely matched the Big Four auditees are to their non-Big Four counterparts in respect of the individual covariates. Panel B of Table 6 reveals that the two samples are also very similar in respect of each of the individual characteristics (variables) in the auditor choice and fee equations (Model 4). Indeed, whilst there are substantial (and significant) differences between the characteristics of the Big Four and non-Big Four auditees before matching (see Table 2), after matching (using the finest calliper of 0.0001) the differences in the means become small and are all statistically insignificant.[31]

Since our sample includes a large number of smaller auditees (due to our relatively low cut-off of £1,000 for total assets), we also conducted additional propensity score matching analysis on samples restricted first to companies with total assets (*TA*) in excess of £10,000 and second to those with *TA* over £1m. Using the 0.0001 calliper, we found that for the first sample (i.e. where *TA* > £10,000), the premium was estimated at 0.2311 (26.0%) and for the second sample (where *TA* > £1m) at 0.1753 (19.2%). Both estimates were statistically significant and even although the latter is somewhat lower than the estimated premium for our main sample, it is within the range reported by Moizer (1997). Furthermore, using the wider calliper of 0.001 on both these size-restricted samples and on two separate sub-samples of firms with *TA* above and below the median (for the full sample) yielded very similar results. Finally, we estimated the models without the instrument (*CHTA*) and our findings were unchanged.[32]

On the basis of matched samples that are very similar in terms of their observed characteristics, therefore, we find strong evidence of a Big Four premium of a similar magnitude to that found in studies employing OLS. Unlike those provided by the Heckman approach, these estimates are not sensitive to model specification and do not impose assumptions of linearity. Moreover, variation in

[27] The conditional mean independence assumption is that the choice of regime (Big Four auditee or non-Big Four auditee) is not dependent on the regime once the matching variables (*Z*) are taken into account. This means, in practice, that the values of *Z* should not depend on the type of regime. We therefore use all the regressors in the fees equation as matching instruments (*Z*) and make the reasonable assumption that all the measured characteristics are pre-determined before the choice of auditor is made.

[28] Leuven and Sianesi (2003) provide details of the propensity score matching module (psmatch2) for use with Stata statistical software.

[29] The analysis in Table 6 is based on differences in the log audit fees to allow comparison with previous findings. We also conducted this analysis using untransformed audit fees and obtained very similar results.

[30] The results based on the 0.001 and 0.0001 callipers in columns 4 and 5 of Table 6 are conducted without replacement, though they are virtually unchanged when we do allow replacement, with a mean difference in log audit fees between Big Four and non-Big Four auditees of 0.2566 and 0.2639 for the 0.001 and 0.0001 callipers respectively (both estimates were statistically significant).

[31] We find similar (unreported) results using the 0.001 calliper, where the only variable approaching statistical significance is *SQSUBS* (difference in means of $-0.06$; $p = 0.066$).

[32] These additional results are unreported for brevity but are available from the authors upon request.

**Table 6**
**Propensity score matched results†**

**Panel A: Alternative propensity score matching estimators**

| | *Nearest neighbour (with replacement)* | *Nearest neighbour (no replacement)* | *Calliper of 0.001Ψ (no replacement)* | *Calliper of 0.0001 (no replacement)* |
|---|---|---|---|---|
| Mean difference in ln*AFEE* | 0.2642 | 0.3082 | 0.2531 | 0.2613 |
| Big Four premium‡ | 30.2% | 36.1% | 28.8% | 29.9% |
| z-statistic | 8.23** | 15.58** | 12.54** | 10.91** |
| *N*§ | 6076 | 6076 | 5586 | 4814 |
| Mean difference in p-score†† | 0.0003 | 0.2393 | 0.0001 | 0.0000 |
| Min. difference in p-score | 0.0000 | 0.0000 | 0.0000 | 0.0000 |
| Max. difference in p-score | 0.3960 | 0.6848 | 0.0010 | 0.0001 |

**Panel B: Effects of matching (calliper 0.0001) on covariate and propensity score means$^{\varphi}$**

| *Variable* | *Big Four pre-match* | *Non-Big Four pre-match* | *Big Four post-match* | *Non-Big Four post-match* |
|---|---|---|---|---|
| P-SCORE$^{\tau}$ | 0.2246 | 0.0700 | 0.1582 | 0.1582 |
| ln*SAL* | 15.666 | 13.385 | 15.128 | 15.149 |
| ln*TA* | 15.488 | 12.860 | 14.906 | 14.884 |
| *CHTA* | 5.2616 | 0.4891 | 1.3583 | 1.2821 |
| *SQSUBS* | 1.1714 | 0.2999 | 0.7827 | 0.8236 |
| *EXPSAL* | 0.7871 | 0.0235 | 0.0571 | 0.0553 |
| *QUALIF* | 0.0375 | 0.0315 | 0.0420 | 0.0403 |
| *PBAL* | 0.1208 | 0.0380 | 0.0835 | 0.0827 |
| *CONLIAB* | 0.2749 | 0.0967 | 0.2194 | 0.2231 |
| *EXITEM* | 0.5662 | 0.3206 | 0.5131 | 0.5172 |
| *TLTA* | 0.8433 | 0.7731 | 0.7817 | 0.7493 |
| *RTA* | 0.0087 | 0.2279 | 0.0206 | 0.0218 |
| *LOND* | 0.2094 | 0.3426 | 0.2185 | 0.2019 |
| *BUSY* | 0.5583 | 0.4399 | 0.5231 | 0.5330 |

*Notes*

† The probit selection model from which propensity scores (p-scores) are derived is reported in Table 3 (Model 4).

†† p-scores are propensity scores.

‡ Results are based on bootstrapped standard errors (50 replications). The premium is the difference between the mean ln*AFEE* for Big Four auditors and the mean ln*AFEE* for matched Big Four counterparts.

§ Note that each method results in an equal number of matched Big Four and non-Big Four auditees.

Ψ The calliper is the maximum permitted absolute difference in propensity score between matched observations.

** represents statistical significance at the 0.01 level.

$\varphi$ None of the means in Panel B differed significantly after matching.

$\tau$ Note that the mean ($p$) of the predicted probabilities derived from the probit model for the whole sample always equals the prior probability of selection into the unity value of the binary dependent variable; in our case $p = 0.083$ (the proportion of Big Four auditees in the sample).

**Table 7**
**Pre-processed portfolio matched regression results**

| | *Mean* | *Std. dev.* | *Median* | *Min* | *Max* |
|---|---|---|---|---|---|
| *BIG4* coefficient | 0.2724 | 0.0114 | 0.2724 | 0.2369 | 0.3091 |
| Big Four premium | 31.3% | – | 31.3% | 26.7% | 36.2% |
| *BIG4* $t$-statistic | 13.73 | 0.5652 | 13.73 | 11.98 | 15.51 |
| $F$-Value | 832.58 | 30.15 | 833.24 | 717.57 | 926.04 |
| Adjusted $R^2$ | 0.7679 | 0.0040 | 0.7679 | 0.7544 | 0.7792 |

*Notes*
† This table reports the distribution of the *BIG4* coefficient estimate from 2,000 iterations, where each iteration involves a regression (see Model 1 in Table 3) using a total sample of 3,656 companies (i.e. 1,828 Big Four and 1,828 non-Big Four auditees), together with associated $t$-statistics and model $F$ and $R^2$. The total sample of 36,674 companies is first divided on the basis of sales (40 quantiles), exports to sales (11 quantiles), return on total assets (40 quantiles) and the number of subsidiaries (10 quantiles). Each Big Four auditee is then matched to a non-Big Four counterpart belonging to the same quantile for each of the four variables. The number of quantiles for the ratio of exports to sales and number of subsidiaries differs due to a large number of zero values for each variable. The $t$-statistic in each iteration uses robust standard errors.

the matching approach (e.g. changing callipers and allowing replacement observations) indicates that, unlike the Heckman approach, the results are not sensitive (in terms of a significant premium) to changes in samples. It should be emphasised, however, that the additional robustness, and the fact that no assumptions are made about functional form in assessing the premium, come at a cost: any unobservable differences between Big Four and non-Big Four clients which systematically influence both auditor choice and audit fees are effectively assumed to be randomly distributed across the samples of Big Four and non-Big Four clients.

*Pre-processed OLS results*

With our pre-processed analysis, using Stata statistical software, we first partition our sample of 36,674 companies into quantiles on the basis of their actual size, risk and complexity, as these factors have been found to be particularly important determinants of both audit fees and auditor selection (e.g. see Simunic and Stein, 1996; Chaney et al., 2004). We created 40 equally sized quantiles based on sales (*SAL*), 40 equally sized quantiles based on return on total assets (*RTA*), 10 quantiles based on the number of subsidiaries (*SUBS*) and 11 quantiles based on the ratio of exports to sales (*EXPSAL*). We then partition the above quantile samples into companies audited by Big Four and non-Big Four auditors and matched them, so that each individual Big Four auditee had an individual non-Big Four counterpart with concurrent membership of the same size quantiles for *SAL*, *RTA*, *SUBS* and *EXPSAL*. Hence observations are matched where they exhibit (jointly) similar size, risk and complexity characteristics. Note that this pre-processing method can be a more demanding process than propensity score matching since the former matches on the basis of actual values for the four control variables *simultaneously*, rather than on one composite score; that is, each Big Four firm has a non-Big Four counterpart with similar observed size *and* risk *and* complexity characteristics.

A dilemma associated with this matching process is that many Big Four clients have a number of non-Big Four counterparts of similar size, risk and complexity. In order to circumvent this problem, we randomly selected (with replacement) one match for each Big Four auditee. We then combined the non-Big Four auditee sample and the Big Four auditee sample and re-estimated our standard regression equation (Model 1 in Table 3). We repeated this process 2,000 times and obtained a distribution of *BIG4* regression coefficients and their associated robust (White's corrected) $t$-statistics (White, 1980). Each iteration involved samples of 1,828 Big Four auditees and 1,828 matched non-Big Four auditee counterparts (i.e. a total sample in each regression of 3,656). Additional analysis (unreported but available on request) showed that the two (Big Four and non-Big Four) auditee samples were very closely matched on the four matching variables with, on average, all differences being highly statistically insignificant.[33]

[33] For each iteration, we collected the $p$-value and $t$-statistic for mean differences between the Big Four and non-Big Four samples for the four matching variables and out of the 2,000 iterations, there were no significant differences (at $p < 0.05$) in the variables on which we matched. More specifically, the range and mean for the absolute $t$-statistics, respectively, of the four variables were 0.31–0.56 and mean of 0.44 for ln*SAL*; 0.17–1.90 and mean of 1.07 for *RTA*; 0.18–0.51 and mean of 0.34 for *SQ-SUBS*; and 0.02–0.04 with a mean of 0.04 for *EXPSAL*. We also repeated this procedure by controlling for both sales and total assets (together with *SUBS*, *EXPSAL* and *RTA*) and obtained similar results (though inevitably on a smaller sample).

Descriptive statistics from the 2,000 regressions for the *BIG4* coefficient and robust *t*-statistics are reported in Table 7. The table shows that in every case, the *BIG4* coefficient was statistically significant and positive, ranging from 0.2724 (26.7%) to 0.3091 (36.2%), with the mean and median taking the same value of 0.2724 (31.3%). The distribution of *t*-statistics (which are based on White's corrected standard errors) reveals that the *BIG4* coefficient is consistently significant at $p < 0.01$. It is also interesting to note that the range of coefficients implies a premium between 27% and 36%, which is broadly in line with that (16–37%) found in prior literature and of a similar magnitude to our propensity score estimates. Hence, the Big Four premium is persistent after matching on key auditee attributes (size, risk and complexity) and controlling for any remaining confounding influences via OLS regression.

## 5. Conclusions

Since the seminal paper by Simunic (1980), a large number of studies have predicted and found that large auditors have commanded a premium for their services, possibly due to superior audit quality, 'deep pockets' and other reputational effects. Important innovations in the literature by (inter alia) Ireland and Lennox (2002) and Chaney et al. (2004) challenged findings based on OLS regressions. The latter paper overturned much of the prior research by stating that, given their firm specific characteristics, private UK companies that chose a Big Five auditor would have paid more had they chosen a non-Big Five auditor, thus leading to the conclusion that the large auditor premium does not exist and that the audit market is properly organised (*ibid.*: 70). Although this is an economically persuasive conclusion, the econometrics literature suggests that selection effects estimated using the Heckman procedure may be highly sensitive to model specification and collinearity; consequently, standard OLS regression can produce more accurate estimates than their two-stage counterparts (Hartman, 1991; Stolzenberg and Relles, 1997).

Applying the Oaxaca-Blinder decomposition to OLS regression results, we find that the majority of the Big Four premium is attributable to large differences in the characteristics of Big Four and non-Big Four auditees. Using this approach, we estimate the Big Four premium at 29–31%, using the estimates of linear equations for samples with markedly different characteristics. An important question, however, is whether the premium correctly reflects what the change in fees would be if auditees changed auditor type. The OB decomposition assumes each auditee can change auditor (unconstrained) to the other type and pay the corresponding counterfactual fees.[34] Our results based on the Heckman correction (which allows for unobservable differences between Big Four and non-Big Four auditees) confirm prior reports in other areas of social science research that this procedure does not represent a panacea for estimating selection effects. The standard Heckman two-step method is highly sensitive to model specification, collinearity and to sample composition. The results also appear highly dependent on the estimation technique adopted. A comparison of the standard Heckman two-step and maximum likelihood estimators revealed large differences in the estimate of the selection effect.

Our findings relating to the impact of correctly identifying the Heckman model with an appropriate instrument are of key importance, since the difficulty in obtaining such an exclusion variable is often a major barrier to valid implementation of the method (see, e.g. Bryson et al., 2002); and ideally, the instrument should *not* be strongly correlated with the remaining selection model regressors. Furthermore, although we present evidence of a significant selection effect, this does not support the interpretation of cost minimisation and does suggest that Big Four auditors receive higher audit fees. The potentially critical finding of Chaney et al. (2004) that private firms select the type of auditor that provides the cheapest service once unobservable auditee characteristics are taken into account is therefore unsupported by our analysis of private UK independent firms.

Whilst two-stage estimators allow for potential unobserved characteristics, our results demonstrate that the advantages of these methods are unproven, and should be traded off against their sensitivity. Three different studies of UK private companies have produced three different sets of results relating to the impact of unobservable variables on premiums using the Heckman estimator, but all reported significant premiums using standard models. Although it is possible (but unlikely) that unobservable factors vary systematically between these studies, further research seems warranted into the causes of this variation (e.g. whether unobservables are non-linear in form, model identification and explanatory variables utilised), and the nature of any such characteristics.

When we employ a more stable matching approach to estimate the premium by comparing the audit fees for Big Four and non-Big Four auditees of a similar degree of size, risk and complexity, we find a persistent premium of a magnitude in line with that found in prior single-stage OLS audit fee studies. The propensity score matching and pre-processed matched regression estimates of the Big

[34] We noted in the paper, for example, that this is highly improbable in respect of the largest auditees (e.g. FTSE 100 companies), which are *de facto* constrained to Big Four auditors. Linear extrapolation outside the common support region in such cases may lead to inaccurate premium estimates; and here propensity score and pre-processed portfolio matching approaches may provide more accurate estimates.

Four premium are around 30%, with such matching estimators being highly robust to changes in model specification, and uninfluenced by assumptions of functional form. Although extant analytical research suggests unobservable factors such as insider knowledge of future cash flows may be important in determining auditor choice, there is disagreement on the magnitude and direction of these effects (cf. Titman and Trueman, 1986; Datar et al., 1991). Further research is warranted on the identification and implementation of empirical proxies for these characteristics (including variables capturing the presence and efficiency of the internal audit function and financial information and control systems), and the inclusion of such variables in more robust matching analyses. For example, Collis et al. (2004) find, using questionnaire data, that demand for voluntary audit of small companies is related to the perceived improvements in internal controls and in the quality of information provided (e.g. to banks). Such methods may provide insights into differential demand for Big Four and non-Big Four audits although may themselves be subject to measurement bias. Taken together with the findings of previous research that large auditors produce higher quality audits (e.g. Blokdijk et al., 2006), and with the findings of a contemporaneous study by Francis and Lennox (2008), our results suggest that the Big Four premium persists and that Heckman-based research that the premium vanishes once selection is allowed for should be treated with caution.

## Appendix

### *(i) Oaxaca-Blinder decomposition*

Standard pooled OLS regression using a Big Four dummy indicator variable in a single equation assumes that all slope coefficients for the Big Four and non-Big Four are identical (i.e.using the notation from equations (1) and (2) in the text, $\alpha_k = \beta_k$, all $k$). The Oaxaca-Blinder decomposition is based on separate regressions for Big Four and non-Big Four clients and emphasises that the observed actual difference in audit fees can partly be attributed to the different characteristics of the two types of auditees and partly to any Big Four premium. Hence, it expresses the difference in the means of audit fees as:

$$\overline{\ln F_{BIG4}} - \overline{\ln F_{NON}} = \sum_{k=1}^{K} a_k(\overline{X_{kBIG4}} - \overline{X_{kNON}}) + b + \sum_{k=1}^{K}(b_k - a_k)\overline{X_{k\,BIG4}} = EXPLAINED_{BIG4} + P_{BIG4} \quad \text{(A1)}$$

$$\overline{\ln F_{BIG4}} - \overline{\ln F_{NON}} = \sum_{k=1}^{K} b_k(\overline{X_{kBIG4}} - \overline{X_{kNON}}) + b + \sum_{k=1}^{K}(b_k - a_k)\overline{X_{k\,NON}} = EXPLAINED_{NON} + P_{NON} \quad \text{(A2)}$$

### *(ii) Heckman selection model*

The Heckman procedure involves estimating the selection term $\lambda$ by modelling the auditor choice process in the first step. Here, each company has an unobserved propensity ($D^*$) to choose a Big Four auditor. $D^*$ is a linear function of $M$ regressors ($Z_m$ $m = 1,.., M$) and other unobservable characteristics ($\varepsilon_{SEL}$). The model can therefore be represented as follows:

$$D^* = \delta + \sum_{m=1}^{M} \delta_m Z_m + \varepsilon_{SEL} \qquad \text{Auditor choice equation} \quad \text{(A3)}$$

$$\ln F_{BIG4} = \alpha + \beta + \sum_{k=1}^{K} \beta_k X_{k\,BIG4} + \varepsilon_{BIG4} \qquad \text{For Big Four auditees} \quad \text{(A4)}$$

$$\ln F_{NON} = \alpha + \beta + \sum_{k=1}^{K} \beta_k X_{k\,NON} + \varepsilon_{NON} \qquad \text{For non-Big Four auditees} \quad \text{(A5)}$$

If $D^* > 0$, $D = 1$ and we observe $\ln F = \ln F_{BIG4}$. Otherwise $D^* \leq 0$, $D = 0$ and $\ln F = \ln F_{NON}$. The model assumes that the errors of the selection and fee equations are jointly normal with zero means, constant variances and covariances: $E(\varepsilon_{SEL}\varepsilon_{NON}) = \sigma_{SNON}$ and $E(\varepsilon_{SEL}\varepsilon_{BIG4}) = \sigma_{SBIG4}$. The implied correlation between the unobservable factors determining the choice of auditor type and audit fees, together with the assumption of normality, enable the estimation of this model.[35] The importance of the normality assumption is discussed by Greene (2003: 789)

[35] For instance, companies more likely to employ Big Four auditors (i.e. have 'large' $\varepsilon_{SEL}$) given their observable characteristics ($Z$) are likely to value unobservable aspects of Big Four auditors' services more highly (i.e. have 'large' $\varepsilon_{BIG4}$).

and by van der Klaauw and Koning (2003); with the latter reporting that coefficient estimates do not appear sensitive to departures from normality. The model may be estimated directly by maximum likelihood but it is more common to employ the Heckman two-step method. This involves the estimation of an augmented regression equation with necessary adjustments to the formulae for the standard errors:

$$\ln F_{BIG4} = \alpha + \beta + \sum_{k=1}^{K} \beta_k X_{k\,BIG4} + \sigma_{SBIG4}\lambda_{BIG4} + \upsilon \qquad \text{For Big Four auditees} \qquad \text{(A6)}$$

$$\ln F_{NON} = \alpha + \sum_{k=1}^{K} \alpha_k X_{k\,NON} - \sigma_{SNON}\lambda_{NON} + \upsilon \qquad \text{For non-Big Four auditees} \qquad \text{(A7)}$$

$$\text{where } \lambda_{BIG4} = \frac{\phi(\sum_{m=1}^{M}\delta_m Z_m)}{\Phi(\sum_{m=1}^{M}\delta_m Z_m)} \text{ and } \lambda_{NON} = \frac{\phi(-\sum_{m=1}^{M}\delta_m Z_m)}{\Phi(-\sum_{m=1}^{M}\delta_m Z_m)} \qquad \text{(A8)}$$

and $\phi$ is the normal density function and $\Phi$ the normal distribution function. Note that the fee equation for non-Big Four auditees is estimated with selection into non-Big Four (i.e. the dependent variable for the probit is $ND = 1$ if the firm is a non-Big Four auditee). The coefficient of the selection term in this estimation is therefore the covariance between the error in the selection equation determining whether $ND = 1$ and $\varepsilon_{NON}$, i.e. an estimate of $-\sigma_{SNON}$. In the interests of comparability between the two equations, all results in the paper for the non-Big Four fee equation report estimates of $\sigma_{SNON}$.

### *(iii) Matching estimators*

Each Big Four auditee can be defined in terms of the values of some observable variables ($\mathbf{Z} \equiv \{Z_1,..,Z_M\}$). Matching uses $\mathbf{Z}$ to produce a comparable non-Big Four auditee for each Big Four auditee. Let $M$ be the set of $N_M$ matched pairs of firms. The estimated treatment effect for each matched Big Four auditee is:

$$D(Z_i) = \ln F_{BIG4}(Z_i) - \ln F_{NON}(Z_i) \qquad i \in M \qquad \text{(A9)}$$

The estimated Big Four premium ($\Delta$), or the treatment effect, is the sample mean of these differences across all values of $Z$ in $M$ or the difference in the sample means. Hence:

$$\Delta = \left[\frac{1}{N_M}\sum_{i\in M} \ln F_{BIG4}(Z_i)\right] - \left[\frac{1}{N_M}\sum_{i\in M} \ln F_{NON}(Z_i)\right] \qquad \text{(A10)}$$

$$\Delta = \overline{\ln F_{BIG4M}} - \overline{\ln F_{NONM}} \qquad \text{(A11)}$$

where the subscript $M$ indicates that the mean is for companies in the matched sample.

Matching relies heavily on two assumptions: the conditional independence assumption (CIA) and common support. The CIA requires the value of audit fees to be independent of auditor type given the values of the observable variables.[36] To illustrate, assume that auditee size is the only determinant of audit fees. Consider a simple comparison of the mean audit fees paid by all Big Four auditees with those paid by all non-Big Four auditees, thus treating auditee size as an unobservable variable. This only makes sense if auditee size is not a determinant of auditor choice. However, if we compare samples of Big Four and non-Big Four auditees with the same size distribution we can ignore the effect of auditee size. CIA states formally that any unobservable variation in fees after adjusting for auditee size has the same random distribution for each type of auditee. Such adjustments are only possible if Big Four and non-Big Four auditees exist of comparable size.

The common support assumption merely states that such comparable auditees exist. For example, if all very large auditees were only audited by the Big Four, then it would be impossible to find comparable non-Big Four auditees. If the Big Four premium is regarded as payable by any auditee, matching assumptions imply that auditees are freely able to switch auditor and incur the corresponding counterfactual fees. It is therefore assumed that any systematic effect of the choice of auditor ($D$) on audit fees can be entirely explained in terms of some observable variables ($\mathbf{Z}$). In practice, $\mathbf{Z}$ is interpreted as the set of determinants of the auditor choice decision. Consider the following semi-parametric matching model (with the selection equation repeated for reference):

[36] More formally: $\ln F_{BIG4}, \ln F_{NON} \perp D \mid \mathbf{Z}$

$$D^* = \delta + \sum_{m=1}^{M} \delta_m Z_m + \varepsilon_{SEL} \qquad \text{Auditor choice equation} \qquad (A12)$$

$$\ln F_{BIG4} = \mu_{BIG4}(Z) + \varepsilon_{BIG4} \qquad \text{For Big Four auditees} \qquad (A13)$$

$$\ln F_{NON} = \mu_{NON}(Z) + \varepsilon_{NON} \qquad \text{For non-Big Four auditees} \qquad (A14)$$

If $D^* > 0, D = 1$ and we observe $\ln F = \ln F_{BIG4}$. Otherwise $D^* \leq 0, D = 0$ and $\ln F = \ln F_{NON}$. The audit fee equations above have additive errors but may be nonlinear in the conditional mean values ($\mu_{BIG4}$ and $\mu_{NON}$). These means may depend on other regressors but the CIA means that the only relevant determinants are contained in ($\mathbf{Z}$). It also means that differences in unobservable auditee characteristics conditional on $\mathbf{Z}$ are random. Matching therefore provides an elegant method of eliminating explained differences. Since it compares samples of auditees with the same characteristics, it is unnecessary to specify the functional form for $\mu_{BIG4}(\mathbf{Z})$ and $\mu_{NON}(\mathbf{Z})$ and to rely on linear projections to produce counterfactuals – an important advantage where a non-linear relationship whose form is often unknown may exist between the dependent variable and the regressors. In practical applications, propensity score matching overcomes the 'curse of dimensionality' problem (which is encountered where $\mathbf{Z}$ contains numerous variables, resulting in a small number of suitable observations for matching purposes) by matching only on one variable (the propensity score), defined as:

$$p(\mathbf{Z}) \equiv \Pr(D = 1 \mid \mathbf{Z}) \qquad (A15)$$

The estimated premium for each matched Big Four auditee is defined as:

$$\Delta(\mathbf{Z}_i) = \ln F_{BIG4}(p(\mathbf{Z}_i)) - \ln F_{NON}(p(\mathbf{Z}_i)) \qquad i \in \Omega \qquad (A16)$$

where $\Omega$ is the set of $N_\Omega$ matched pairs of firms. The estimated Big Four premium ($\Delta$) is the sample mean of these differences across all values of $p(\mathbf{Z})$ in $\Omega$, or the difference in the sample means:

$$\Delta = \left[\frac{1}{N_\Omega}\sum_{i\in\Omega} \ln F_{BIG4}(p(Z_i))\right] - \left[\frac{1}{N_\Omega}\sum_{i\in\Omega} \ln F_{NON}(p(Z_i))\right] \qquad (A17)$$

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# Book Review

**The Routledge Companion to Accounting History.** Editors: *John Richard Edwards and Stephen P. Walker.* Routledge, Taylor & Francis Group, London and New York. 2008. xvii and 619 pp. ISBN 9780415410946. £125.

As stated by its editors, the aim of this volume is 'to offer an introduction to the shifting arenas which attract the attention of accounting historians, relate the findings of their research and address the controversies that energise debate in the field' (p. 1). That aim has been handsomely achieved, and even exceeded.

This comprehensive and up-to-date survey of accounting history comprises an introduction followed by 28 chapters which are grouped into seven parts. The chapters generally have a common form, with each providing an overview of the selected topic, a summary and analysis of relevant research, and the identification of opportunities for further inquiries. At the end of each chapter the references cited are listed, along with four or five 'key works' which are nominated as the starting point for further reading.

The 37 contributors include a wealth of leading scholars, including many whose work has extended beyond accounting history. The contributors are diverse in terms of their research approaches and are drawn from a range of countries. However, as one of the editors, Stephen P. Walker, notes in his opening chapter, the global development of accounting history as an academic discipline has been uneven. Unsurprisingly, this would appear to be particularly the case in terms of English language scholarship, making it perhaps inevitable that contributors from the UK, USA, Canada and Australia predominate. They are, however, supplemented by authors based in Spain, New Zealand, Belgium and Ireland.

Compendium books are prone to having a fragmented quality on account of the different approaches and styles of the contributors, especially when there are 37 of them. That the *Routledge Companion* escapes such a failing is a testimony to some combination of the provision of detailed guidelines to contributors, the willingness of contributors to comply with those guidelines, and thorough and diligent editorial work. The editors, wisely in the opinion of this reviewer, did 'not ... attempt to present a grand or metanarrative from a particular philosophical perspective' (p. 3). The reward for this decision is a volume which acknowledges and embraces the diversity of accounting history scholarship. However, the editors have still very skilfully managed to craft a 'whole' rather than a series of disparate parts. This is evident, for example, in the cross references between chapters and particularly in the editors' introduction which outlines and justifies the structure and purpose of the book. This coherence is also indicative of careful planning by the editors and their judicious selection of topics. Established themes in accounting history research (such as bookkeeping, professionalisation, and the US railroads) are evident, but so too are topic areas that have emerged only more recently (for example, emancipation, religion, and the creative arts). This combination provides for a comprehensive survey which avoids unnecessary overlaps and serves to highlight the still evolving scope of the discipline.

The magnitude of the *Routledge Companion* precludes a detailed chapter-by-chapter commentary here. However, the following references to selected chapters are intended to give some general indication of the book's scope. In Part 1, 'The discipline', Stephen P. Walker has contributed a chapter entitled 'Structures, territories and tribes' which, as well as being of interest to established accounting history researchers, will be essential reading for those wishing to commence research into accounting's past. The chapters in Part 2 dealing with 'Technologies' span ancient accounting through to the impacts of mechanisation and computerisation, with an interleaving chapter by David Oldroyd and Alisdair Dobie on the perennially debated subject of the origins and significance of 'Bookkeeping'. Part 3, on 'Theory and practice', includes standout chapters by Thomas A. Lee on 'Financial accounting theory' and Richard Fleischman on 'Management accounting: theory and practice'. Part 4 deals with 'Institutions', with a chapter by Christopher Poullaos on 'Professionalisation' providing a comprehensive review of the abundant research – now spanning a diverse array of locations – into the occupational status of accounting. 'Economy' is the title of the fifth part, which features a chapter by Thomas A. Lee, Frank L. Clarke and Graeme W. Dean on 'Scandals'. It highlights the potency of learning from the past, even if such lessons are sometimes neglected by regulators and policy makers. Six chapters are devoted to Part 6 on 'Society and

culture' – highlighting the very active recent engagement of accounting historians in this area. Sam McKinistry's chapter on the 'Creative arts' typifies the more novel contexts in which accounting historians are now active. Part 7, 'Polity' contains three chapters ranging across the subjects of the state, the military and taxation.

The *Routledge Companion* is likely to be of interest to a broad range of readers. Established accounting history scholars will benefit from having access to a single volume which identifies and summarises research that has accumulated over time and been published in a variety of media. Those academics responsible for delivering a unit in accounting history within coursework programs will, without reservation, be able to adopt this book as an anchor point for student reading. Commencing accounting history researchers will find within this single volume an accessible yet comprehensive entrée to the discipline, and be stimulated by the abundant and enticing suggestions for further research.

More generally, anyone with an interest in accounting is certain to find within this volume something (and perhaps lots) that is of interest. In Evelyn Waugh's celebrated novel *Brideshead Revisited* one of the characters comments that 'we possess nothing certainly except the past'. Of course, we cannot possess, or know, all of the past and some degree of uncertainty will attach to that which we do know. But all empirical research is essentially history, even if it is not labelled as such, and what accounting is today is the culmination of its past. To try and understand accounting – what it is, its nature and its consequences – requires recourse to its origins, and the *Routledge Companion* provides a comprehensive, accurate and up-to-date overview of what is known about accounting's past. That in itself is a very commendable achievement. However, the significance of this book extends beyond just providing an excellent survey of accounting history research. To their great credit, the editors and contributors have also created a work that testifies to the importance of that research.

University of Ballarat,
Australia Brian West

## For the attention of prospective and recent post-graduates in accounting history

### THE ACADEMY OF ACCOUNTING HISTORIANS 2009 VANGERMEERSCH MANUSCRIPT AWARD

In 1988, The Academy of Accounting Historians established an annual manuscript award to encourage scholars new to the field to pursue historical research. An historical manuscript on any aspect of the field of accounting, broadly defined, is appropriate for submission.

### ELIGIBILITY AND GUIDELINES FOR SUBMISSIONS

Any accounting faculty member, who holds a full-time appointment and who received his/her masters/doctorate within seven years previous to the date of submission, is eligible to be considered for this award. Co-authored manuscripts will be considered (if at least one co-author received his/her master/doctorate within the last seven years). Manuscripts must conform to the style requirements of the *Accounting Historians Journal.* Previously published manuscripts or manuscripts under review are not eligible for consideration.

Each manuscript should be submitted by **30 June 2009** in a Word file as an e-mail attachment to the chair of the Vangermeersch Manuscript Award Committee, Professor Lee Parker (lee.parker@unisa.edu.au).

A cover letter, indicating the author's mailing address, the date of the award of the masters/doctoral degree, and a statement that the manuscript has not been published or is not currently being considered for publication should be included in the submission packet.

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The Vangermeersch Manuscript Award Committee will evaluate submitted manuscripts on a blind-review basis and select one recipient each year. The author will receive a US $1,000 stipend and a plaque to recognise his/her outstanding achievement in historical research. In the case of co-authored manuscripts, only the junior faculty member(s) will receive prizes. The winning manuscript will be published in the *Accounting Historians Journal* after an appropriate review. The award will be given annually unless the Manuscript Award Committee determines that no submission warrants recognition as an outstanding manuscript.

# Guide for Authors

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Manuscripts should be in English and consist of original unpublished work not currently being considered for publication elsewhere. The paper should be submitted electronically as Microsoft Word files via e-mail to abr@cch.co.uk. An electronic acknowledgement of receipt will be sent by return.

If you have any problems sending a Word file as an attachment to an e-mail, please send an e-mail to the above address explaining the difficulty.

Experience has shown that papers that have already benefited from critical comment from colleagues at seminars or at conferences have a much better chance of acceptance. Where the paper shares data with another paper, an electronic copy of the other paper must be provided. Authors of accepted papers will also be asked to assign exclusive copyright to the publishers.

*Presentation*

Each submission should include a cover page in a separate Word file that contains the names, affiliations, and contact details of the author(s). The cover page should include the title of the paper and any acknowledgements to third parties. The main body of the paper should appear in a separate Word file, starting with the title of the paper, but without the author's name, followed by an abstract of 150–200 words. Keywords (maximum of five) should be inserted immediately following the abstract. The main body of the paper should start on the next page. In order to ensure an anonymous review, authors should endeavour to avoid identifying themselves. Section headings should be numbered using Arabic numerals.

*Tables and figures*

Each table and figure should bear an Arabic number and a title and should be referred to in the text. Where tables and figures are supplied in a format that can not be edited within a Word document, delay in publication may result. Sources should be clearly stated. Sufficient details should be provided in the heading and body of each table and figure to reduce to a minimum the need for the cross-referencing by readers to other parts of the manuscript. Tables, diagrams, figures and charts should be included at the end of the manuscript on separate pages, with their position in the main body of the text being indicated.

*Footnotes*

Footnotes should be used only in order to avoid interrupting the continuity of the text, and should not be used to excess. They should be numbered consecutively throughout the manuscript with superscript Arabic numerals. They should not normally be used in book reviews.

*Mathematics*

Authors are asked to use mathematics only if it contributes to the clarity and economy of the article. Where possible, authors should restrict the use of mathematics to an appendix. Equations should be numbered in parentheses, flush with the right hand margin. Authors of mathematically-oriented papers written in Scientific Word or with some other mathematical word processing package are advised to consult with the editor about the format before making a formal submission, in order to avoid technical difficulties later.

*References*

References should be listed at the end of the paper and referred to in the text as, for example (Zeff, 1980: 24). Wherever appropriate, the reference should include a page or chapter number in the book or journal in question. Only works cited in the article should be included in the list. Citations to institutional works should if possible employ acronyms or short titles. If an author's name is mentioned in the text it need not be repeated in the citation, e.g.'Tippett and Whittington (1995: 209) state…'

In the list of references, titles of journals should omit an initial 'The' but should not otherwise be abbreviated. The entries should be arranged in alphabetical order by surname of the first author. Multiple authors should be listed in chronological order of publication, e.g.:

Accounting Standards Steering Committee (1975). *The Corporate Report*. London: ASC.

Tippett, M. and Whittington, G. (1995). 'An empirical evaluation of an induced theory of financial ratios'. *Accounting and Business Research*, 25(3): 208–218.

Watts, R.L. and Zimmerman, J.L. (1986). Positive Accounting Theory. Englewood Cliffs, NJ: Prentice Hall.

*Style and spelling*

Abbreviations of institutional names should be written as, for example, FASB and not F.A.S.B.; those of Latin terms should contain stops (thus i.e. not ie). Words such as 'realise' should be spelt with an 's', not a 'z'. Single quotation marks should be used, not double.